



Parental assortative mating and the intergenerational transmission of human capital [☆]

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ABSTRACT

We study the contribution of parental educational assortative mating to the intergenerational transmission of educational attainment. We develop an empirical model for educational correlations within the family in which parental educational sorting can translate into intergenerational transmission jointly by both parents, or transmission can originate from each parent independently. Estimating the model using educational attainment from Danish population-based administrative data for over 400,000 families, we find that on average 75 percent of the intergenerational correlation in education is driven by the joint contribution of the parents. We also document a 38 percent decline of assortative mating in education for parents born between the early 1920s and the early 1950s. While the raw correlations also show decreases in father- and mother-specific intergenerational transmissions of educational attainment, our model shows that once we decompose all factors of intergenerational mobility, the share of intergenerational transmission accounted for by parent-specific factors increased from 7 to 27 percent; an increase compensated by a corresponding fall in joint intergenerational transmission from both parents, leaving total intergenerational persistence unchanged. The mechanisms of intergenerational transmission have changed, with an increased importance of one-to-one parent-child relationships.

1. Introduction

Understanding the mechanisms of intergenerational transmission of socioeconomic status is key for the design of policies aimed at promoting equal opportunities. Since Becker (1973), parental human capital has been considered to be the engine of intergenerational transmission of advantage.¹ Parental assortative mating in education – the tendency for individuals with the same education to mate with one another more frequently than would be expected under a random mating pattern – can also affect intergenerational transmission by generating segmentation in the distribution of parental human capital. This segmentation results in an increasing share of households characterized by high (or low) levels of human capital for both parents, so parental sorting can act as an amplifier of intergenerational persistence in human capital (Doepke et al., 2019).

In this paper, we study the contribution of parental educational sorting to the intergenerational transmission of human capital. We develop an empirical model of educational correlations within the family in which human capital of parents is generated by factors shared between parents due to assortative mating, some of which can be fully transferred to children, while other factors are not passed on to children. In addition, each parent may transmit human capital independently to their children due to factors that are not shared with the co-parent. Therefore, in our model intergenerational persistence can arise both from the joint contribution of the parents and from the independent contributions of each parent.

We estimate the model using educational attainment from Danish population-based administrative data for over 400,000 families. Identifying the contribution of parental assortative mating to the intergenerational transmission of human capital is achieved by exploiting

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¹ The extensive literature on intergenerational mobility that has emerged is reviewed in Solon (1999), Black and Devereux (2011) and Mogstad and Torsvik (2021).

family quartets composed of two parents and two children. We find that about 75 percent of the intergenerational correlation in education is driven by the *joint contribution* of the two parents, while the *independent transmission* from each parent is only one sixth of the size of the joint parental contribution, with father-specific transmission being larger for sons, and mother-specific transmission being larger for daughters. This pattern suggests that most of the persistence in educational attainment is due to factors that both parents share and transmit to their offspring.

Exploring trends over time, we document that parental assortative mating in Denmark has declined by 38 percent for parents born between 1920 and 1952.² While the raw correlations also show decreases in father- and mother-specific intergenerational transmissions of educational attainment, our model shows that once we decompose all factors of intergenerational mobility, the share of intergenerational transmission accounted for by parent-specific factors increased from 7 to 27 percent. This increase in parent-specific intergenerational transmission reflects a corresponding fall in joint intergenerational transmission from both parents, leaving total intergenerational persistence unchanged. The nature of intergenerational transmission is changing, with a growing importance of one-to-one parent-child relationships.

This finding about the nature of intergenerational transmission can, in part, be explained by the increase in the time parents have been spending with their children over recent decades in Denmark and across other western countries, especially on activities enhancing child human capital potential with an increasing involvement of fathers (e.g., Sandberg and Hofferth, 2001; Sayer et al., 2004; Guryan et al., 2008; Gimenez-Nadal and Sevilla, 2012; Sani and Treas, 2016). The composition of parental time with children has also changed with the emergence of intensive parenting, including activities in which both parents spend time together with their children (joint family time) but also activities in which children are spending time with only one of the parents (Craig et al., 2014).

This paper relates to a growing literature in economics on the relationship between assortative mating and intergenerational transmission. One strand of the literature focuses on educational sorting in the *offspring generation*, taking the view that non-random sorting among offspring might affect measures of intergenerational mobility, slowing down regression to the mean in socioeconomic status between parents and children (e.g., Chadwick and Solon, 2002; Ermisch et al., 2006; Holmlund, 2020).

Another strand of the literature, which is closer to our study, examines the effect of sorting in the *parental generation* on the socioeconomic status of the offspring generation. Although this literature has mostly focused on inequality (e.g., Fernández and Rogerson, 2001; de la Croix and Doepke, 2003; Fernández et al., 2005; Bratsberg et al., 2018), the influence of parental sorting on intergenerational mobility may be even more important than on inequality (e.g., Kremer, 1997). Nevertheless, the empirical evidence on the link between parental sorting and intergenerational mobility is still limited. Among the few existing studies, Güell et al. (2015), exploit the joint distribution of surnames and economic outcomes and provide evidence of a decrease in intergenerational mobility in Catalonia throughout the 20th century, suggesting the increase in assortative mating to be an important driver of this secular fall in intergenerational mobility. Collado et al. (2019), exploit extended families to infer the characteristics of multi-generational transmission

and find a predominant role of assortative mating for explaining intergenerational persistence.

The empirical model of educational correlations within the family that we develop in our study contributes to this literature in three ways. First, we decompose parental assortative mating into intergenerational and intragenerational components, where only the former contributes to parent-child educational correlations. Second, we quantify how much parental assortative mating contributes to intergenerational persistence relative to factors that are passed on to children from each parent independently from the co-parent. Third, we trace out how the components of educational correlations within the family have evolved over time.

The time evolution of the relative importance of joint versus parent-specific contributions is also related to the literature documenting the evolution of assortative mating, which concludes that changes in assortative mating over time have a limited impact on overall income inequality (e.g., Eika et al., 2019; Frémeaux and Lefranc, 2017). We contribute to this literature by showing that the impact of the declining importance of assortative mating for intergenerational transmission observed in Denmark is compensated by countervailing trends in parent-specific transmission of education to children, leaving total intergenerational persistence unchanged.

The approach we propose in this paper is not without limitations. Most importantly, although it is intuitive to allow for joint versus parent-specific transmission of human capital, we cannot identify all the factors that can account for this distinction. Differences in parental time allocation and its evolution over time is one among several possible explanations. The model also focuses on overall trends of assortative mating without allowing for differences in marital sorting across educational groups; trends that have been found to differ both within and between countries. Our proposed model offers a decomposition of intergenerational persistence and assortative mating in education that complements other approaches in the literature.

The remainder of the paper is structured as follows. In the next section we describe the data source and provide a summary of raw statistics on the extent of intra-family correlations in education. In Section 3 we outline the econometric model of educational correlations of family members, while in Section 4 we present the estimation results. In Section 5 we provide a concluding discussion.

2. Data description and raw patterns of educational correlations within the family

We use data from administrative registers of the Danish population. The civil registration system was established in 1968 and everyone resident in Denmark then and since has been issued with a unique personal identification number, which has subsequently been used in all national registers enabling accurate linkage. We base our analysis on *family quartets* composed of two parents and their first two children observed in the civil register.³ Although, in principle, data on siblings are not needed to estimate parental assortative mating or intergenerational correlations separately from one another, we show in the next section that observing two children per family is key for identifying the contribution of parental assortative mating to the intergenerational transmission of human capital.

We match the civil register with data on the educational attainment of family members. We observe qualifications in the education register, and use the norms produced by the Ministry of Education to calculate the time that would be taken to obtain each qualification by the shortest

² Trends in educational assortative mating vary across countries. The declining trend in educational assortative mating we report for Denmark is consistent with the evidence in Eika et al. (2019) and we document this in Section 2. The declining trend of assortative mating in Denmark is similar to the trends reported in Norway (Bratsberg et al., 2018) and Sweden (Holmlund, 2020). For the U.S., Eika et al. (2019) report an increasing trend in educational assortative mating from 1940 until the 1980s, after which it changed little, while Güell et al. (2015) report an increasing trend in educational assortative mating in Catalonia, Spain.

³ In robustness checks we consider alternative sampling schemes, either extending the sample by including half siblings, or by using the second and third children rather than the first and second. All the conclusions reached in the paper are robust to these alternative sampling schemes.

Table 1
Descriptive statistics.

	(1) Number of persons				(2) Years of education	
	By role in the family				Mean	S.D.
	Mother	Father	Child 1	Child 2		
All birth cohorts		417,746			12.30	3.04
Birth cohort						
1921	648	5,171			9.91	3.54
1924	3,033	12,009			9.99	3.53
1927	7,983	20,358			10.27	3.53
1930	17,036	31,801			10.53	3.48
1933	30,931	43,452			10.67	3.39
1936	47,842	52,290			10.73	3.35
1939	57,374	56,593			10.92	3.33
1942	65,529	62,880			11.28	3.27
1945	75,491	68,291			11.69	3.14
1948	64,877	47,036			11.98	2.99
1951	47,002	17,865			12.09	2.80
1957			55,998	8,243	12.86	2.41
1960			61,203	42,109	12.87	2.32
1963			65,902	60,513	13.03	2.25
1966			71,126	67,655	13.22	2.22
1969			62,321	62,802	13.37	2.28
1972			55,689	65,454	13.69	2.29
1975			35,078	56,036	14.02	2.24
1978			10,429	37,077	14.28	2.22
1981			0	14,826	14.38	2.26
1984			0	2,691	14.39	2.32
1987			0	340	14.56	2.28

Note: The left panel of the table reports the number of individuals in the estimating sample, their role in the family and birth cohort, where the cohorts are defined as non-overlapping groups of individuals who are born over three consecutive years. The table indicates each cohort by its central year. The right panel of the table reports summary statistics for the distribution of years of education by birth cohort.

route. Educational attainment is imputed from the qualification with the highest normed duration on October 31 in the year of turning 29.⁴

We draw data on the full population of families resident in Denmark 1980–2017 where we observe at least two children for parents born from 1920 (the first birth year for which we have educational information) to 1979, a total of 639,516 families. We exclude families for which either parent was younger than 18 at the birth of a child, resulting in a sample of 624,883 families. Next, we select families where we observe a child is born in 1956 or later, which is the first year the parent-child linkage is complete; and where the second child is born in 1988 or earlier, which is the last birth year for which we observe educational attainment at age 29, because our data were sourced in early 2018. Selecting on children's year of birth returns a sample of 530,390 families. Next, we drop families with sibling age spacing smaller than one year, or larger than ten years, corresponding to the first and last percentile of the distribution of age spacing, resulting in a sample of 524,577 families. We also drop families whose first child was born after 1978 to allow for a maximum sibling spacing of ten years also among younger families, resulting in a sample of 477,953 families. Among these families, those with parents born in the 1950s tend to be considerably younger at first birth by design, with average age at first birth declining from about 23 for mothers born in 1952, or earlier, to 18 for mothers born in 1960. To avoid selectivity into early parenthood among later cohorts, we therefore further exclude families with either parent born after 1952, yielding a sample of 432,715 families. Finally, we drop families with missing information on educational attainment for any member, resulting in an estimating sample of 417,746 families and 1,670,984 individuals.

We group individuals into three-year cohorts based on birth year. Table 1 describes the cohort structure of the data, indicating each three-year cohort by its central year. The average year of birth is around 1939 for fathers and 1942 for mothers, while it is around 1966 for first chil-

⁴ Educational qualifications obtained before 1971 are self-reports from the 1970 population census. From 1971, educational institutions report qualifications obtained to the Ministry of Education. See Jensen and Rasmussen (2011).

Table 2

Raw correlations in years of education among family members, by gender composition of the children.

	Sibling composition			
	Two boys	Two girls	Mixed	All families
Parents	0.50	0.50	0.50	0.50
Children	0.34	0.34	0.31	0.33
Mother-child	0.28	0.27	0.28	0.28
Father-child	0.29	0.26	0.27	0.28

Note: The table reports empirical correlation coefficients of years of education between family members. Correlations are computed after residualizing years of school on year-of-birth dummies and are estimated for each combination of birth cohorts of the family members, after excluding cells with fewer than 100 pairs. Reported correlations are averaged over birth cohorts.

dren and 1969 for second children. The table also shows the evolution of the distribution of years of education over birth cohorts. In the 66 years that separate the first and the last cohort, average attainment has grown by about 4.5 years, while the standard deviation has decreased by about 1 year. These patterns are in line with expectations given the educational expansions that have characterized the 20th century. To remove common trends in education that may induce spurious patterns of correlation, we residualize years of education on year-of-birth dummies throughout the analysis. Moreover, we conduct the empirical analysis by modelling intra-family educational *correlations* which are insensitive to variations in marginal dispersion.

Table 2 provides a summary of raw intra-family correlations in years of education for different family types defined by the gender mix of the two children. Correlations are estimated by birth cohorts after excluding cells with fewer than 100 pairs.⁵ The educational correlation among

⁵ Estimating correlations by cohorts will enable us to consider secular changes. We also compute bootstrapped standard errors of these correlation coefficients that will be used later in the paper to adjust inference.

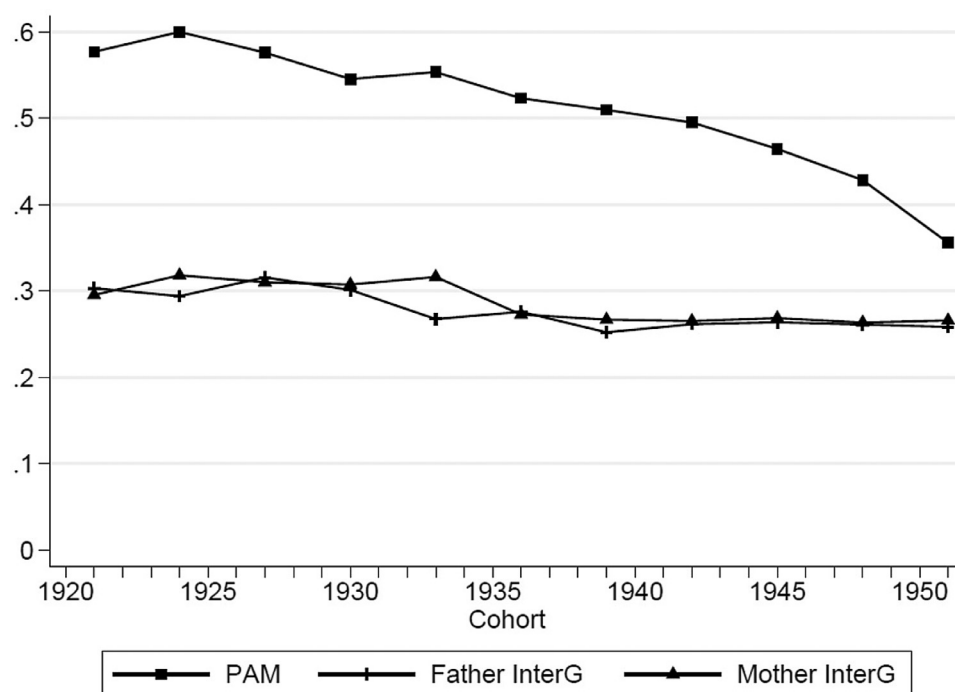


Fig. 1. Parental assortative mating and intergenerational transmission - Raw correlations. Note: The graph plots empirical correlation coefficients of years of education between family members. The line labelled PAM shows the correlation between parents; the line labelled Father InterG shows the father-child correlation; the line labelled Mother InterG shows the mother-child correlation. Correlations are computed after residualizing years of schooling on year-of-birth dummies and are estimated for each combination of birth cohorts of the family members, after excluding cells with fewer than 100 pairs. Correlations are plotted by birth cohort of the oldest member in the pair.

parents is 0.5, which is not only substantial but also stable across family types. This correlation is our basic raw measure of parental assortative mating. The intergenerational correlation in education is substantially lower at about 0.27, exhibiting almost no variation across family types for mothers, while there is some evidence that it is larger in all-son families for fathers. Comparing these intergenerational correlations with those for parental assortative mating suggests that not all factors that make parents similar to each other are transmitted to their children, because otherwise intergenerational and spousal correlations would be of similar sizes. Table 2 also shows that the sibling correlation in education is sensitive to the type of family, being larger in families with same-gender children than it is in those with mixed-gender children. This heterogeneity of the sibling correlation in education by the gender mix of the children is consistent with evidence of a greater exposure to shared community effects for siblings of the same gender (e.g., Martin and Fabes, 2001; Fabes et al., 2003).

Taking advantage of the cohort structure of the estimated raw correlations, we plot them against cohort of birth, after pooling across family types. Fig. 1 shows that parental educational assortative mating has declined by more than a third for the cohorts of parents from 1921 to 1951, which corresponds approximately with matches formed from the mid-1940s to the late 1970s.⁶ This pattern is consistent with the declining trend of educational assortative mating reported by Eika et al. (2019) for

Denmark between 1980 and 2013.⁷ The declining trend of assortative mating in Denmark is also consistent with the trends reported in other Scandinavian countries (Bratsberg et al., 2018, and Holmlund, 2020). In Fig. 1 we also observe 13 percent declines of both father- and mother-specific intergenerational correlations in education, which change from about 0.3 to 0.26 over thirty years.

That the declining assortative mating of the parents is reflected in intergenerational transmission only to a very limited extent is consistent with two possible explanations that are not mutually exclusive. One explanation may be that the *joint contribution* of parents to the human capital of their children has been declining, but this decline has been compensated by the *individual contributions* of each parent and their ability to independently transmit human capital to their children. For example, parenting style might have been shifting to a one-to-one parent-child relationship, which is consistent with the increase in the time parents spend with their children in Denmark and in other western countries (e.g., Guryan et al., 2008; Sani and Treas, 2016) and the emergence of intensive parenting (Craig et al., 2014). An alternative explanation

⁶ We also find a broadly similar decline in assortative mating as in Fig. 1 when we compute the Spearman correlation coefficient using four educational groups (i.e., less than High School/High School/Some College/College). Furthermore, to check whether this measure of assortative mating is sensitive to the number of groups as suggested by Gihleb and Lang (2020), we also used a five-group partition – distinguishing within the “less than High School” between those who only completed compulsory education and those who enrolled for one or two years of non-compulsory education – obtaining again similar results. Also if we extend the period of observation of parental assortative mating to the cohorts born until 1970, corresponding to matches formed until the mid to late 1990s, we find that the decline in assortative mating continues with spousal correlation in education halving with respect to the initial value in Fig. 1. However, these younger cohorts do not contribute to our estimating sample because their children are too young to observe completed education.

⁷ Eika et al. (2019) report a measure of assortative mating which is the odds ratio between the actual joint probability that husband and wife have the same level of education and the probability of that event under random matching. This measure yields a set of marital sorting estimates by level of education, that can be aggregated into an overall weighted average of assortative mating. For Denmark, Eika et al. (2019) use married couples aged 26–60 between 1980 and 2013 and provide estimates of assortative mating by educational level, but not their weighted average. After replicating their estimates by educational level, we find that the weighted average declines from 1.75 in 1980 to 1.55 in 2013. Applying this measure of assortative mating on our cohort-wise data, we find a decline from 1.71 for the 1921 cohort to 1.51 for the 1951 cohort; this decline is 28 percent of the decline that would be needed to reach random matching (corresponding to a marital sorting parameter equal to 1). The measure of assortative mating reported in Fig. 1 based on the correlation coefficient of spouses' years of education declines from 0.57 for the 1921 cohort to 0.35 for the 1951 cohort; this decline is 38 percent of the decline that would be needed to reach random matching (corresponding to a correlation coefficient equal to 0). Using the same data and sample design as in Eika et al. (2019), we also compute for Denmark the correlation coefficient of spouses' years of education finding a decline from 0.5 in 1980 to 0.45 in 2013.

for the decline of assortative mating may be that over time there is less sorting based on traits that are irrelevant for intergenerational transmission. This alternative explanation indicates a more efficient matching process in the marriage market, in which traits that do not contribute to the human capital of children are less and less valued. Assessing the relevance of the two alternative scenarios is not possible based only on the contrast between raw parental assortative mating and intergenerational transmission. The model presented in the next Section allows us to understand their relative importance.

3. An econometric model of educational correlations among family members

In this section, we present a model for the educational attainment of family members which we use to measure the contribution of parental assortative mating to the intergenerational transmission of human capital. Families are composed of four members: father (F), mother (M), first child (C_1) and second child (C_2). Let y_{ij} denote years of education for person i in family j residualized on year-of-birth dummies to remove common trends, and let $H(i)$ denote the role of i in the family, that is $H(i) \in \{M, F, C_1, C_2\}$. For the parental generation, years of education of parent i are given by

$$y_{ij} = a_{ij} + \gamma_{Aj} + \gamma_{H(i)j} + \mu_j, \quad H(i) = F, M \quad (1)$$

that is, they are factored into the sum of four orthogonal zero mean components. The first term, a_{ij} , represents sources of human capital, such as idiosyncratic ability or luck, that are specific to parent i and are not shared with other family members.

The second and third terms denote human capital generating factors that are passed across generations. We distinguish two such factors. The first factor, denoted as γ_{Aj} , is shared between parents and represents the contribution of parental assortative mating to intergenerational transmission of human capital. This term captures attitudes or values shared between parents that are relevant for the acquisition of education and are transmitted jointly by both parents to children. For example, both parents share values regarding educational attainment or parenting style, skills and attitudes, such as the importance of hard work, which they jointly transfer to children. The second factor of intergenerational transmission denoted as $\gamma_{H(i)j}$, is parent-specific and, therefore, it is not shared between parents. This factor captures specific skills, attitudes, and values that each parent transmits to children independently from the co-parent.

Finally, the fourth term μ_j , represents determinants of human capital that are shared between parents, but are not passed on to the children, i.e., a purely intragenerational component of parental assortative mating. For example, parents may share traits, habits or behaviors that can hamper human capital (say smoking) that they do not want to pass to the children. More generally, there may be frictions in the transmission process such that not all factors shared by the parents are passed on to the next generation.

In the offspring generation, years of education of child i are given by

$$y_{ij} = a_{ij} + \gamma_{Aj} + \gamma_{Mj} + \gamma_{Fj} + \gamma_{Mj}\gamma_{Fj} + \theta_j, \quad H(i) = C_1, C_2 \quad (2)$$

that is, they are factored into the sum of the following orthogonal components: (i) idiosyncratic factors, a_{ij} , such as ability and luck, that are specific to child i and are not shared with other family members; (ii) all factors that are received from the parent generation (i.e. the shared factors – γ_{Aj} – the parent-specific factors from the mother – γ_{Mj} – and the father – γ_{Fj} – and their interaction – $\gamma_{Mj}\gamma_{Fj}$ – to allow for complementarity between the parent-specific components); and (iii) other intra-generational influences, θ_j , such as school or community factors that siblings share independently from their parents.

By assuming orthogonality across the components of Eq. (2), and in particular between each of the γ terms and θ , we are effectively

ruling out sorting of families across schools and communities. In Bingley et al. (2021) we show how sorting can be identified in models of the sibling correlation by exploiting the timing of families' mobility across communities; results for Denmark indicate that community-related effects tend to be overestimated when ignoring sorting. Therefore, the correlation of the intra-generational term θ_j in Eq. (2) must be seen as an upper bound on the relevance of these influences that siblings share independently from their parents within the overall sibling correlation of education.

We assume that each of the zero mean human capital generating factors described above has variance σ_s^2 with $s \in \{a, \gamma_A, \gamma_F, \gamma_M, \mu, \theta\}$. We estimate the contributions of these factors to educational correlations within the family by matching moment restrictions generated by the model to the empirical correlations estimated from the data. Because these correlations capture similarities in education between family members, they will be functions only of the variances of shared factors and not of idiosyncratic factors, and only the former can be estimated from the empirical correlations.

Starting from the parents, their correlation of educational attainment implied by the model is given by

$$\rho_{MF} = \sigma_{\gamma_A}^2 + \sigma_{\mu}^2. \quad (3)$$

The parameter ρ_{MF} is our measure of parental assortative mating and is the sum of the variances of two factors: those that are shared between parents and are transmitted to the children ($\sigma_{\gamma_A}^2$), and those that are shared between parents but are not transmitted to the children (σ_{μ}^2). In contrast, parent-specific determinants of the intergenerational transmission of human capital do not contribute to parental assortative mating.

Looking across generations, the intergenerational correlation of education between each parent and each child is given by

$$\rho_{PC} = \sigma_{\gamma_A}^2 + \sigma_{\gamma_P}^2, \quad P = M, F \quad (4)$$

which is the sum of the variances of the factors that are shared between parents and are transmitted to the children ($\sigma_{\gamma_A}^2$), and the parent-specific factors transmitted by each parent to each of the children ($\sigma_{\gamma_P}^2$) independently from the co-parent. Note that the three parameters of intergenerational transmission $\sigma_{\gamma_A}^2$, $\sigma_{\gamma_M}^2$ and $\sigma_{\gamma_F}^2$ can also be used to derive a measure of total intergenerational correlation, that is the one that originates from all the factors that the parents pass on to each child, either jointly or individually: $\rho_{MFC} = \sigma_{\gamma_A}^2 + \sigma_{\gamma_M}^2 + \sigma_{\gamma_F}^2 + \sigma_{\gamma_M}^2\sigma_{\gamma_F}^2$, where the last term is the variance of $\gamma_{Mj}\gamma_{Fj}$, which equals the product of the variances because γ_{Mj} and γ_{Fj} are zero mean independent random variables. This correlation has no empirical counterpart and therefore is not a moment that we can use in estimation, but is a statistic that we can compute post-estimation.

The sibling correlation, based on the model assumptions, is given by

$$\rho_{C_1C_2} = \sigma_{\gamma_A}^2 + \sigma_{\gamma_M}^2 + \sigma_{\gamma_F}^2 + \sigma_{\gamma_M}^2\sigma_{\gamma_F}^2 + \sigma_{\theta}^2, \quad (5)$$

which is the sum of the variances of all the factors transmitted intergenerationally (i.e., the total intergenerational correlation ρ_{MFC}) and the intragenerational factors that siblings share independently from the parents (σ_{θ}^2).

It is important to emphasize that this is a system of four equations (two intragenerational equations – (3) and (5) – and two intergenerational equations – (4), one for each parent) in five parameters ($\sigma_{\gamma_A}^2$, $\sigma_{\gamma_M}^2$, $\sigma_{\gamma_F}^2$, σ_{μ}^2 , σ_{θ}^2), and it is not identified. For identification, therefore, we need an additional restriction.

To solve the indeterminacy, we exploit variation in the gender composition of the children and restrict mixed-gender children to share only factors stemming from the family environment, while same-gender children share influences originating both from within the family as well as outside the family. That is, we restrict

$$\theta_j = 0 \text{ if } g_{ij} \neq g_{i'j}; \quad H(i) = C_1, \quad H(i') = C_2 \quad (6)$$

where g is a gender indicator. The rationale for this restriction is that, particularly at young ages, interactions independent from the family environment occur predominantly among same-gender peers (e.g., Martin and Fabes, 2001; Fabes et al., 2003). The evidence shown in Section 2 of lower educational correlations for mixed-gender children, compared to those of the same gender, provides some corroboration suggesting that shared influences among same-gender children may represent an upper bound on the values of the parameter for mixed-gender children. Under restriction (6), the difference in educational correlations between same-gender and mixed-gender children identifies the parameter σ_{θ}^2 . To operationalize this idea, we compute the empirical correlations by family type based on the gender composition of the children (mixed, all sons, all daughters) and use the empirical correlations of all family types jointly in estimation.

We estimate the model by Minimum Distance and match the empirical educational correlations with their counterparts predicted by the model, using bootstrapped standard errors of the correlations to weight the minimization problem. The correlation function implied by the model for members H and H' is given by:

$$\begin{aligned} f(\omega)_{HH'} = & \sigma_{\gamma A}^2 \left[I(H = M)I(H' = F) + I(H = P)I(H' = C) + \right. \\ & \left. I(H = C_1)I(H' = C_2) \right] \\ & + \sigma_{\mu}^2 I(H = M)I(H' = F) \\ & + \sigma_{\gamma M}^2 [I(H = M)I(H' = C) + I(H = C_1)I(H' = C_2)] \\ & + \sigma_{\gamma F}^2 [I(H = F)I(H' = C) + I(H = C_1)I(H' = C_2)] \\ & + \sigma_{\gamma M}^2 \sigma_{\gamma F}^2 I(H = C_1)I(H' = C_2) \\ & + \sigma_{\theta}^2 I(H = C_1)I(H' = C_2)I(g_H = g_{H'}), \end{aligned} \quad (7)$$

where $I(\cdot)$ is an indicator function, ω is the vector collecting all model parameters, $P \in \{M, F\}$, $C \in \{C_1, C_2\}$, and the regression does not include a constant such that estimated parameters can be interpreted as levels (of the components) of correlation coefficients.

4. Results

We start the discussion of results by providing some benchmark estimates from intergenerational OLS regressions of years of education in Table 3. We consider both children of each family in each regression and cluster standard errors at the family level. All variables are measured as deviations from generational means, so the regressions do not include a constant. In Column (1), the regression coefficient of child years of education on father's years of education is equal to 0.2, indicating that an additional year of father's education is associated with an increase of child's education of about two and a half months. Rescaling the regression coefficient with the ratio of the standard deviations of years of education for fathers and children (which is equal to 1.4), the implied intergenerational correlation of years of education is 0.28. This is very close to the estimate of Table 2 obtained by averaging correlations across cohorts, after excluding cells with fewer than 100 pairs. Similarly, in Column (2), the regression coefficient for mother's education is equal to 0.21 and the implied intergenerational correlation is equal to 0.29.

As noted by Holmlund et al. (2011), regressing child outcomes on each parent's outcome in isolation from their spouse, delivers estimates

that are a convolution of the direct effect of that parent and the impact of parental assortative mating. As a way of netting out the latter, they suggest including the outcomes of both parents in the intergenerational regressions. We present estimates from this specification in Column (3) of Table 3, finding that the direct effect for each parent is indeed lower (by about a third) relative to the corresponding coefficients in Columns 1 and 2. These estimates suggest that assortative mating indeed plays a role in shaping intergenerational transmission; however, they do not provide any measure of this effect, which is instead the aim of the model we propose and estimate in this paper.

In a final set of benchmark estimates, we extend the OLS specification by including the interaction between father's and mother's education in order to capture any complementarity of parental inputs. Column (4) of Table 3 shows the regression coefficient on the interaction is positive and statistically significant, which suggests that there is complementarity, although its magnitude is negligible compared with the baseline effects.

In what follows, we report and discuss Minimum Distance estimates for the baseline model in Section 4.1, by gender of the children in Section 4.2 and over time in Section 4.3.

4.1. Baseline estimates

The baseline estimates of Panel (A) in Table 4 suggest a correlation of educational attainment between parents of about 0.48, which is the sum of the component of parental assortative mating that is transmitted intergenerationally to the children ($\sigma_{\gamma A}^2 = 0.2403$, s.e. = 0.0048) and the component that is not transmitted and instead remains within the generation of the parents ($\sigma_{\mu}^2 = 0.2418$, s.e. = 0.0054). That is, on average across birth cohorts and family types, about half of the factors shared by parents are transmitted intergenerationally.

Panel (A) of Table 4 also reports the parent-specific factors ($\sigma_{\gamma M}^2$, $\sigma_{\gamma F}^2$) that each parent transmits independently to their children. Combining the joint ($\sigma_{\gamma A}^2$) with the parent-specific factors, we obtain an intergenerational correlation of education which is equal to 0.2755 for mothers (0.2403 + 0.0352) and 0.2790 for fathers (0.2403 + 0.0387). Thus, parent-specific contributions to intergenerational transmission appear to be of modest size, each about one sixth of the size of the joint parental contribution, and very similar for mothers and fathers. The complementarity effect is given by the product of parent-specific factors and is equal to 0.0013 (s.e. = 0.0003), which is negligible compared to the joint and parent-specific contributions.

With these estimates we can compute the *total* intergenerational correlation, which measures transmission from both parents (either joint or parent-specific) to a child; as shown in Section 3, this total correlation is not directly observed in the data, but it is estimated by the model. Total intergenerational correlation is equal to 0.3155 (= 0.2403 + 0.0352 + 0.0387 + 0.0013). This sum implies that about 75 percent of the intergenerational correlation in education is driven by the joint parental contribution (0.2403/0.3155), suggesting a major role of parental assortative mating in accounting for intergenerational transmission. Our model – based on the direct estimation of the share of parental assortative mating within the intergenerational correlation –

Table 3
OLS intergenerational regressions.

	(1)		(2)		(3)		(4)	
	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Years of education of								
Father	0.201	0.001			0.129	0.001	0.130	0.001
Mother			0.217	0.001	0.153	0.001	0.149	0.001
Father × Mother							0.010	0.0003

Note: The table reports coefficients from OLS regression of child years of education on parental years of education with robust standard errors clustered at the level of the family. Each regression is run on 835,492 observations.

Table 4
Parameter estimates of educational correlations model.

	Coeff.	S.E
(A) Baseline		
Intragenerational Parental Assortative Mating σ_{μ}^2	0.2418	0.0054
Intergenerational Parental Assortative Mating $\sigma_{\gamma A}^2$	0.2403	0.0048
Intergenerational Mother-specific $\sigma_{\gamma M}^2$	0.0352	0.0043
Intergenerational Father-specific $\sigma_{\gamma F}^2$	0.0387	0.0043
Residual sibling σ_{θ}^2	0.0431	0.0058
(B) Heterogeneity by family type		
Intragenerational Parental Assortative Mating σ_{μ}^2	0.2421	0.0065
Intergenerational Parental Assortative Mating $\sigma_{\gamma A, MX}^2$	0.2414	0.0055
Intergenerational Parental Assortative Mating $\sigma_{\gamma A, BB}^2$	0.2396	0.0081
Intergenerational Parental Assortative Mating $\sigma_{\gamma A, GG}^2$	0.2376	0.0083
Intergenerational Mother-specific $\sigma_{\gamma M, MX}^2$	0.0357	0.0047
Intergenerational Mother-specific $\sigma_{\gamma M, BB}^2$	0.0316	0.0090
Intergenerational Mother-specific $\sigma_{\gamma M, GG}^2$	0.0392	0.0093
Intergenerational Father-specific $\sigma_{\gamma F, MX}^2$	0.0370	0.0047
Intergenerational Father-specific $\sigma_{\gamma F, BB}^2$	0.0556	0.0090
Intergenerational Father-specific $\sigma_{\gamma F, GG}^2$	0.0237	0.0093
Residual sibling $\sigma_{\theta, BB}^2$	0.0298	0.0119
Residual sibling $\sigma_{\theta, GG}^2$	0.0575	0.0122
(C) Cohort trends		
Intragenerational Parental Assortative Mating σ_{μ}^2 - Intercept (c. 1921)	0.3496	0.0103
Intragenerational Parental Assortative Mating σ_{μ}^2 - Cohort trend	-0.0058	0.0005
Intergenerational Parental Assortative Mating $\sigma_{\gamma A}^2$ - Intercept (c.1921)	0.2947	0.0078
Intergenerational Parental Assortative Mating $\sigma_{\gamma A}^2$ - Cohort trend	-0.0022	0.0004
Intergenerational Mother-specific $\sigma_{\gamma M}^2$ - Intercept (c. 1921)	0.0086	0.0056
Intergenerational Mother-specific $\sigma_{\gamma M}^2$ - Cohort trend	0.0009	0.0003
Intergenerational Father-specific- Intercept (c. 1921)	0.0124	0.0065
Intergenerational Father-specific- Cohort trend	0.0007	0.0003
Residual sibling σ_{θ}^2 - Intercept (c. 1957)	0.0431	0.0070
Residual sibling σ_{θ}^2 - Cohort trend	0.0001	0.0006

Note: The table reports parameter estimates obtained by Minimum Distance. Robust standard errors are obtained by weighting the minimization problem with the inverse of the bootstrapped variance of empirical moments. In the model with heterogeneity by family types in Panel (B), MX denotes families with mixed-gender siblings, BB denotes families with two sons and GG denotes families with two daughters. The number of persons is equal to 1,670,984 and the number of moments is equal to 1,297.

points to a greater role of parental assortative mating compared with the one implicit in the indirect assessment of Table 3 based on the contrast of OLS regressions with different conditioning sets.

Finally, Panel (A) of Table 4 reports the estimate of the intragenerational factors that siblings share independently from their parents ($\sigma_{\theta}^2 = 0.0431$, s.e. = 0.0050), which is small and accounts for about 12 percent of the overall sibling correlation. The latter is equal to 0.3586 (calculated as the sum of all factors transmitted intergenerationally – shared – $\sigma_{\gamma A}^2$ – and parent-specific – $\sigma_{\gamma M}^2$, $\sigma_{\gamma F}^2$ and their product – plus the intragenerational factors that children share independently from their parents – σ_{θ}^2).⁸ As all these estimates are derived under the restriction that mixed-gender children share only influences within the family, the relatively small estimates of the residual sibling effect σ_{θ}^2 from same-gender children – that can be thought of as upper bounds for the mixed-gender estimates – suggest that the assumption of zero independent correlation for mixed-gender children is reasonable.

4.2. Heterogeneity of educational correlations

The finding that, compared to the joint contributions, the independent parental contributions to intergenerational transmission are of sec-

ond order importance, may miss part of the parental effect if each parent's ability to influence their children's educational attainment depends on the family environment. For example, fathers could have greater influence in a mostly male environment, while mothers in a mostly female environment. To investigate whether child gender matters for the magnitude of each parent's independent influences, we extend the model by allowing parameters to differ according to the gender mix of the children.

In particular, we interact family type with the three intergenerational parameters $\sigma_{\gamma A}^2$, $\sigma_{\gamma M}^2$ and $\sigma_{\gamma F}^2$, which are the ones needed to decompose the intergenerational correlation into parental assortative mating and parent-specific transmission and are central to our research question. We also interact family type with the residual sibling correlation σ_{θ}^2 , which in the baseline model is effectively already family-type-specific, being zero for families with mixed-gender children. It is worth noting that we cannot interact family type with *all* the parameters of the baseline model because a saturated interaction would re-introduce the under-identification problem (discussed in Section 3) within each family type. To avoid this issue, we restrict the intragenerational assortative mating parameter σ_{μ}^2 to be constant across family types.

The results, reported in Panel B of Table 4, show evidence that gender matters for intergenerational transmission. For mothers, we find that the mother-specific contribution to the intergenerational correlation is about a quarter larger when both children are daughters ($\sigma_{\gamma M, GG}^2$), compared to when both children are sons ($\sigma_{\gamma M, BB}^2$), but the difference is not statistically significant. For fathers, we find that the father-specific contribution to the intergenerational correlation is about twice as large when both children are sons ($\sigma_{\gamma F, BB}^2$) rather than daughters ($\sigma_{\gamma F, GG}^2$), while father-specific transmission with mixed-gender children ($\sigma_{\gamma F, MX}^2$) is in-between the same-gender estimates, and all differences are statis-

⁸ Using a model of sibling correlations and intergenerational transmission from the father only, Bingley and Cappellari (2019) obtain that the share of brothers' correlation in education accounted for by factors that the siblings share independently of the father is about 26 percent (see their Table 4). Because that model does not consider mothers, its residual sibling correlation includes the mother-specific intergenerational transmission and the complementarity between parent-specific components. The sum of mother-specific intergenerational transmission, complementarity and residual sibling correlation in our model accounts for about 24 percent of the total sibling correlation.

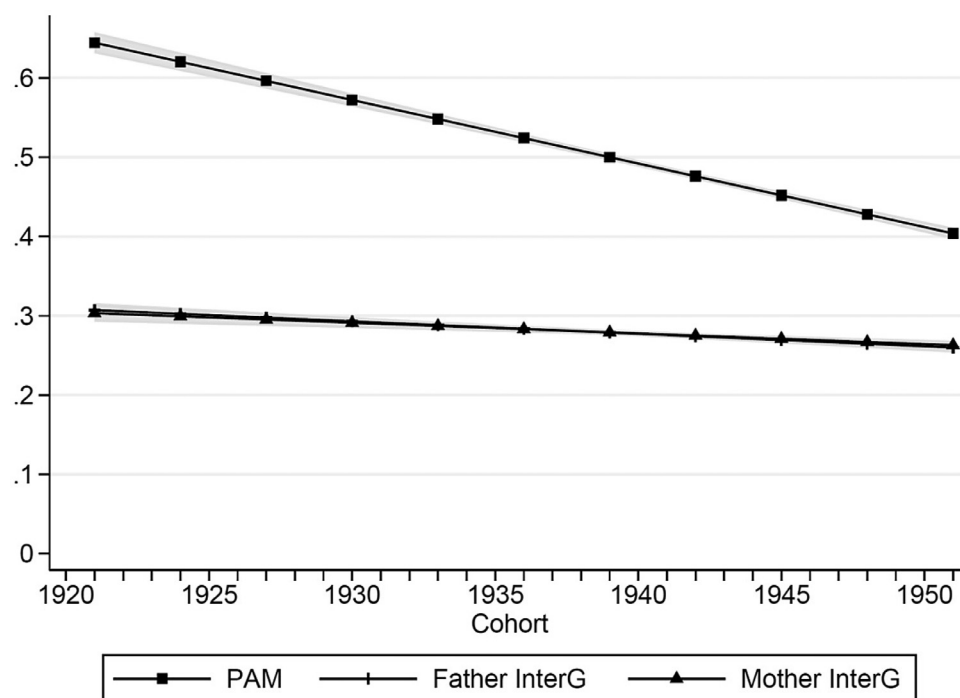


Fig. 2. Parental assortative mating and intergenerational transmission - Predicted correlations.

Note: The graph plots predicted correlation coefficients of years of education between family members. The shaded areas are the 95% confidence intervals of the predictions. The line labelled PAM shows the correlation between parents; the line labelled Father InterG shows the father-child correlation; the line labelled Mother InterG shows the mother-child correlation. Correlations are estimated from the model with cohort trends of Section 4.3.

tically significant (at the 10 percent level for mixed vs. all-daughters, or lower). Together with the finding that the intergenerational component of assortative mating is rather stable across family types, these estimates show only modest differences in parental assortative mating between family types in shaping *total* intergenerational transmission, which varies between 72 percent for families with all sons and 78 percent for all daughters.

The remaining parameter that varies with family type in Panel B of Table 4 is the residual sibling correlation σ_{θ}^2 , which is larger for daughters than for sons ($\sigma_{\theta,GG}^2 = 0.057$ vs. $\sigma_{\theta,BB}^2 = 0.030$), and the difference is significant at the 5 percent level.

4.3. Cohort analysis

We extend the analysis by exploiting the cohort structure of the empirical moments to estimate the evolution in the components of assortative mating and intergenerational transmission over time. As reported in Fig. 1, there is a marked decline for Denmark in parental educational assortative mating between the cohorts 1921 and 1951, while the evolution of intergenerational educational correlations is also declining, albeit at a more modest pace. As discussed in Section 2, the different trends in assortative mating and in intergenerational transmission are compatible with two possible explanations. According to the first explanation, there is a decline in the joint intergenerational transmission from both parents, which is mostly compensated by an increase in parent-specific intergenerational transmission. According to the second explanation, the decline in parental assortative mating is only driven by a decline over time in the component of parental assortative mating that is not transmitted from parents to children.

By exploiting the cohort structure within our model, we are able to offer new insights regarding the evolution of parental assortative mating and intergenerational transmission over time. In particular, we interact each of the model parameters with a linear trend in the birth cohort. For all parameters that involve the parents, that is, the parameters of parental assortative mating and intergenerational transmission, the trend is fit using the cohort of the elder parent. For child correlations, we use the cohort of the elder sibling. We fit simple linear trends which, as Fig. 1 suggests, should not be overly restrictive.

The estimates are reported in Panel C of Table 4 and are summarized in Figs. 2 to 4. Fig. 2, which is the counterpart of Fig. 1 predicted from the model, displays an overall marked decline of parental educational assortative mating (the annual change of spousal correlation implied by the estimates is equal to -0.008 – s.e. = 0.0003) and only a minor reduction over time of intergenerational transmission from each parent (the annual change in intergenerational correlations implied by the estimates is -0.0016 for fathers and -0.0013 for mothers – s.e. = 0.0002 in both cases). The parent-specific estimates are very close to each other and not statistically different, resulting in closely overlapping plots in Fig. 2. Fig. 3 decomposes the trend of parental assortative mating into its intra- and inter-generational components, showing that the sharp observed decline of parental educational assortative mating over time is due to both components, with a prevalence of the intragenerational component (in Table 4 the annual change of spousal correlation due to the latter is -0.0057 , versus a change of -0.0022 stemming from the inter-generational component of assortative mating).

Fig. 4 decomposes *total* intergenerational transmission (the sum of all intergenerational factors) into the part due to assortative mating and the part due to parent-specific transmission, showing that, while the total effect is stable, its components have followed quite distinct trends. The contribution of parental assortative mating has declined from 0.29 for the 1921 cohort to 0.23 for the 1951 cohort. In contrast, there is an increase of the parent-specific components. The sum of all the parent-specific factors reported in Fig. 4, that is the father- and mother-specific components and their complementarity, increase from 0.021 in the 1921 cohort to 0.083 in the 1951 cohort, suggesting an increase in the share of intergenerational transmission accounted for by parent-specific factors from 7 to 27 percent.⁹ These findings suggest a change in the nature of intergenerational transmission over time, with an increasing importance of one-to-one parent-child relationships.

⁹ Estimating the time trends of intergenerational transmission by OLS, we found an increase of the mother-specific coefficient over time and of the father-mother interaction, while we found the father-specific coefficient declines over time.

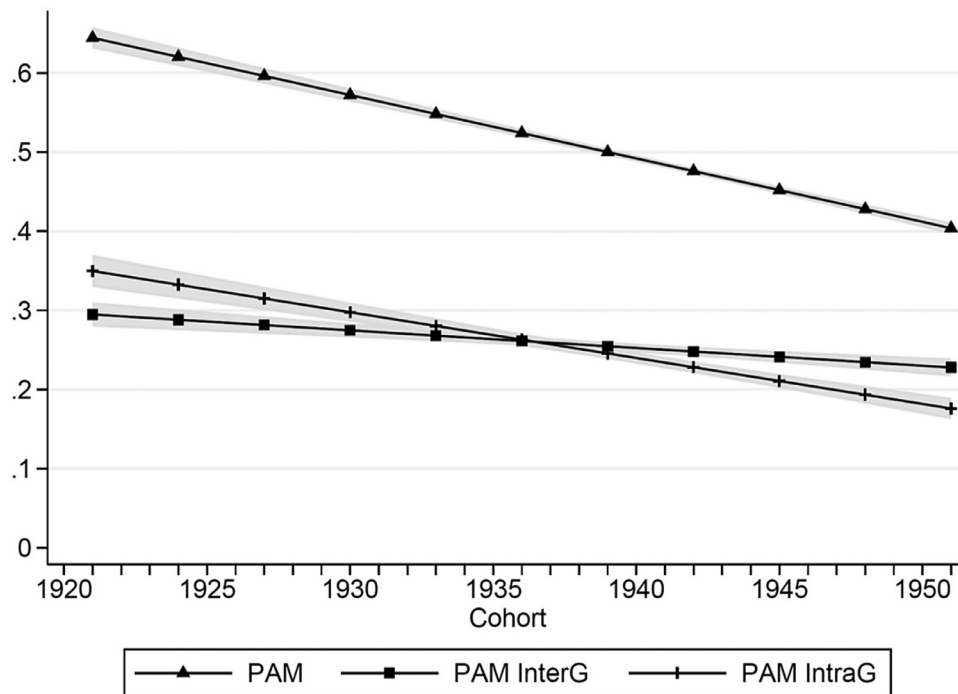


Fig. 3. Predicted decomposition of parental assortative mating.

Note: The graph plots predicted correlation coefficients of years of education between parents and their decomposition into the factors that are transmitted to children and the factors that are not transmitted. The shaded areas are the 95% confidence intervals of the predictions. The line labelled PAM shows the correlation between parents; the line labelled PAM InterG shows the factors that are transmitted to children; the line labelled PAM IntraG shows the factors that are not transmitted. Correlations are estimated from the model with cohort trends of Section 4.3.

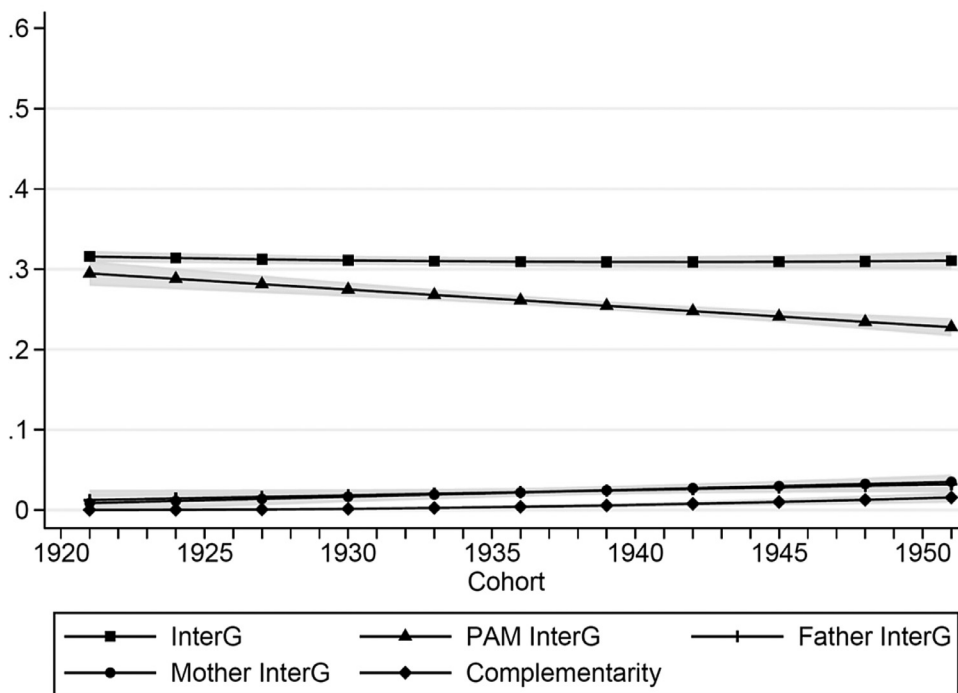


Fig. 4. Predicted decomposition of intergenerational transmission.

Note: The graph plots predicted correlation coefficients of years of education between parents and children and their decomposition into the factors that are transmitted jointly by both parents and the factors that are transmitted individually by each parent. The shaded areas are the 95% confidence intervals of the predictions. The line labelled InterG shows the total intergenerational correlation resulting from the sum of all the factors that are passed from parents to children; the line labelled PAM InterG shows the factors that are transmitted jointly by the parents; the line labelled Father InterG shows the factors that are transmitted only by the father; the line labelled Mother InterG shows the factors that are transmitted only by the mother. Mother InterG line obscures Father InterG because they are so close. The line labelled Complementarity shows the product of the father and mother-specific factors. Correlations are estimated from the model with cohort trends of Section 4.3.

5. Conclusion

We study the importance of parental educational assortative mating (sorting) in shaping the intergenerational persistence of education by transmitting factors shared by both parents to the children, relative to influences transmitted by each parent separately. We develop an empirical model of educational correlations within families composed of two parents and two children. We estimate the model using educational attainment from Danish population-based administrative data for over 400,000 families.

We find that about 75 percent of the intergenerational correlation in education is driven by the joint contribution of the parents, which

points towards a strong role of shared parental inputs in producing the human capital of children. Parent-specific contributions to intergenerational transmission of education are each about a sixth of the size of the joint parental contribution, with father-specific transmission being larger for sons, and mother-specific transmission being larger for daughters.

Considering the evolution of parental assortative mating and intergenerational transmission of educational attainment over time, we document a 38 percent secular decline of parental educational assortative mating for parents born between the early 1920s and the early 1950s, as well as a decrease in the raw intergenerational educational correlations for both fathers and mothers. Our model shows that once we decom-

pose all factors of intergenerational mobility, the share of intergenerational transmission accounted for by parent-specific factors increased from 7 to 27 percent. This increase in parent-specific intergenerational transmission is compensated by a corresponding decline in the share accounted for by the joint transmission from both parents, leaving intergenerational persistence unchanged. This shift from joint to parent-specific transmission implies a reduction in the importance of parental assortative mating for intergenerational transmission over time, with an increase in the importance of one-to-one parent-child relationships, which is consistent with recent trends of increasing parental time devoted to children and the emergence of intensive parenting.

The declining trend of the parental educational assortative mating we report for Denmark, although similar to the evolution in other Scandinavian countries, differs from the trends in other western countries, which also exhibit more variation in marital sorting across educational groups. Understanding these differences in the trends both across and within countries is an interesting avenue for future research.

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