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PhD Thesis

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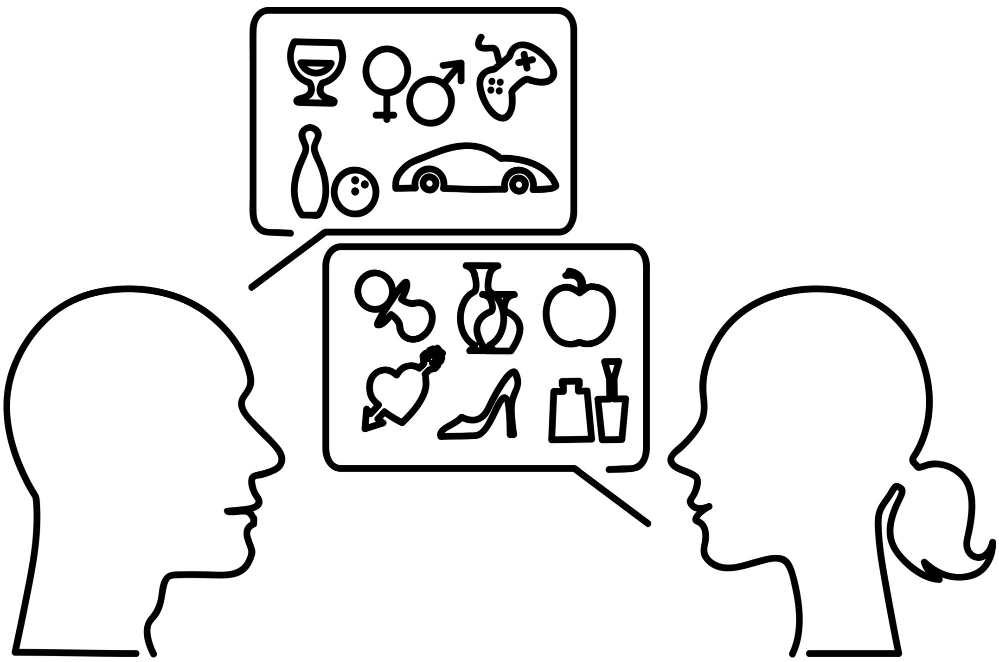
Causes and Consequences of Gender Differences in Human Capital Formation



Supervisor: Mette Gørtz
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Anne Ardila Brenøe

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Summary

The human capital of an individual is an important predictor of social and economic success in life and differs greatly across individuals and groups. Children from disadvantaged socio-economic groups, for instance, typically score lower in cognitive tests than children from more advantaged groups. And women are, on average, less willing to compete and less (over)confident than men. Differences in the formation of human capital across different groups foster behavioral and economic inequality in society. Each individual accumulates, develops, and forms their human capital throughout life. In particular, conditions, interventions, and exposures during early life and childhood are important for human capital development. In this thesis, I focus on causes and consequences of *gender* differences in the accumulation of human capital. Specifically, I examine how different aspects of the social environment during childhood affect gender differences in human capital formation.

This thesis consists of four self-contained chapters, with overlapping themes within the topic of gender and human capital formation within the broader literature in labor economics. Common for all chapters is that I use quasi-experimental empirical strategies with Danish administrative data to provide credible causal estimates; Chapters 1 and 4 further complement the administrative data sets with survey data. While Chapter 2 studies the role of the school environment, Chapters 1, 3, and 4 studies different facets of the childhood family environment. Chapters 1 and 2 study the choice of field of education and occupation, thereby focusing on the *type* of human capital. In contrast, Chapter 3 concerns the length of educational attainment and adult labor market outcomes and Chapter 4 considers time discounting preferences.

Chapter 1 is entitled "Origins of Gender Norms: Sibling Gender Composition and Women's Choice of Occupation and Partner". In this chapter, I examine how one central aspect of the childhood family environment–sibling gender composition– affects women's gender identity, measured through their occupational and partner choice. I causally estimate the effect of having a second-born brother relative to a sister for first-born women. The results show that women with a brother acquire more traditional gender norms with negative consequences for their labor earnings. I provide evidence of increased gender-specialized parenting in families with mixed sex children, suggesting a stronger transmission of traditional gender norms. Finally, I find indications of persistent effects to the next generation of girls.

Chapter 2, "Exposure to More Female Peers Widens the Gender Gap in STEM Participation", is joint work with Ulf Zölitz. We investigates how high school gender composition affects students' participation in STEM college studies. Using Danish administrative data, we exploit idiosyncratic within-school variation in gender composition. We find that having a larger proportion of female peers reduces women's probability of enrolling in and graduating from STEM programs. Men's STEM participation increases with more female peers present. In the long run, women exposed to more female peers earn less because they (1) are less likely to work in STEM occupations, and (2) have more children. Our findings show that the school peer environment has lasting effects on occupational sorting and the gender wage gap.

Chapter 3, "Gender Gaps in the Effects of Childhood Family Environment: Do They Persist into Adulthood?", is joint work with Shelly Lundberg and is forthcoming in *European Economic Review*. We examine the differential effects of family disadvantage on the education and adult labor market outcomes of men and women born between 1966 and 1995. We link parental education and family structure during childhood to male-female and brother-sister differences in adolescent outcomes, educational attainment, and adult earnings and employment. Our results are consistent with U.S. findings that boys benefit more from an advantageous family environment than do girls in terms of grade-school outcomes. Father's education, which has not been examined in previous studies, is particularly important for sons. However, we find a very different pattern of parental influence on adult outcomes. Gender gaps in educational attainment, employment, and earnings are increasing in maternal education, benefiting daughters. Paternal education decreases the gender gaps in educational attainment (favoring sons) and labor market outcomes (favoring daughters). We conclude that differences in the behavior of school-aged boys and girls are poor proxies for differences in skills that drive longer-term outcomes.

Chapter 4, "The Intergenerational Transmission of Time Preferences Persists Across Four Decades", is joint work with Thomas Epper. We study the intergenerational transmission of time preferences, using an experimentally validated survey measure. Parents' and children's impatience is measured four decades apart, thereby eliminating concerns regarding reverse causality. Our results show a substantial transmission of impatience from parents to children. This correlation is insensitive to the inclusion of comprehensive sets of administratively reported controls. We further show that mothers differentially and more strongly transmit impatience to daughters relative to sons, even when comparing siblings. This suggests that nurture affects children's impatience. Finally, the strength of the transmission does not diminish as children age, emphasizing the persistence of preference propagation.

Resumé (Danish Summary)

Et individs humankapital spiller en afgørende rolle for social og økonomisk succes livet og varierer meget på tværs af enkeltpersoner og grupper. Børn fra dårligt stillede socioøkonomiske kår klarer sig typisk dårligere i kognitive test end børn fra bedrestillede kår. Og kvinder er eksempelvis i gennemsnit mindre villige til at konkurrere og mindre overmodige end mænd. Forskelle i dannelsen af humankapital på tværs af forskellige grupper fremmer adfærdsmæssig og økonomisk ulighed i samfundet. Hvert individ akkumulerer, udvikler og danner sin egen humankapital gennem livet. Særligt spiller forhold, interventioner og eksponeringer igennem barndommen en vigtig rolle for udviklingen af humankapital. I denne afhandling fokuserer jeg på årsager til og konsekvenser af kønsforskelle i akkumuleringen af humankapital. Jeg undersøger mere præcist, hvordan forskellige aspekter af det sociale miljø i barndommen påvirker kønsforskelle i dannelsen af humankapital.

Denne afhandling består af fire selvstændige kapitler med overlappende temaer inden for emnet køn og opbygning af humankapital inden for den bredere litteratur inden for arbejdsmarkeds-økonomi. Fælles for alle kapitler er, at jeg bruger kvasi-eksperimentelle empiriske strategier med dansk registerdata for at kunne sige noget om årsagssammenhængene; Kapitel 1 og 4 supplerer yderligere det administrative datasæt med data fra to spørgeskemaundersøgelser. Mens Kapitel 2 undersøger skolemiljøets rolle, belyser Kapitel 1, 3 og 4 forskellige aspekter af barndomsfamilie-miljøet. Kapitel 1 og 2 undersøger valget af uddannelses- og beskæftigelsesområde og belyser derved *typen* af humankapital. I modsætning hertil omhandler Kapitel 3 længden af uddannelse og arbejdsmarkedsudfald i voksenlivet og Kapitel 4 omhandler tidsdiskonteringspræferencer.

Kapitel 1 har titlen “Origins of Gender Norms: Sibling Gender Composition and Women’s Choice of Occupation and Partner”. I dette kapitel undersøger jeg, hvordan et centralt aspekt af barndomsfamiliemiljøet–søskendes kønssammensætning–påvirker kvinders kønsidentitet, målt via deres beskæftigelses- og partnervalg. Jeg estimerer den kausale effekt af at have en lillebror i forhold til en lillesøster for førstefødte kvinder. Resultaterne viser, at kvinder med en bror erhverver mere traditionelle kønsnormer med negative konsekvenser for deres arbejdsindkomst. Jeg viser desuden, at forældre i familier med børn af begge køn øger deres kønsspecialisering, hvilket indikerer en stærkere overførsel af traditionelle kønsnormer i denne type familier. Endelig finder jeg tegn på vedholdende virkninger for den næste generation af piger.

Kapitel 2, “Exposure to More Female Peers Widens the Gender Gap in STEM Participation”, er udarbejdet i samarbejde med Ulf Zölitz. Vi undersøger, hvordan kønssammensætningen i gymnasiet påvirker elevernes valg af universitetsstudie inden for natur-, teknologi-, ingeniør- og matematikvidenskab (forkortet *STEM* på engelsk). Vi udnytter variationen i kønssammensætningen på tværs af årgange inden for det samme gymnasium. Vi finder, at kvinder med en større andel af kvindelige klassekammerater har en lavere sandsynlighed for at påbegynde og færdiggøre studier inden for STEM. Mænds STEM-deltagelse øges, når de har haft en større andel kvindelige klassekammerater. På længere sigt tjener kvinder, der udsættes for flere kvindelige klassekammerater, mindre, fordi de (1) er mindre tilbøjelige til at arbejde i STEM-erhverv og (2) har flere børn. Vores resultater viser, at skolemiljøet har en varig indvirkning på kønssegregering på arbejdsmarkedet og lønforskellen mellem mænd og kvinder.

Kapitel 3, “Gender Gaps in the Effects of Childhood Family Environment: Do They Persist into Adulthood?”, er udarbejdet i samarbejde med Shelly Lundberg og udkommer i *European Economic Review*. Vi undersøger, hvordan mænd og kvinder påvirkes forskelligt af deres barndomsfamiliemiljø med hensyn til færdiggjort uddannelse og arbejdsmarkedsudfald for årgangene født mellem 1966 og 1995. Vi sammenligner forskelle mellem mænd og kvinders og mellem brødre og søstres humankapital afhængigt af deres forældres uddannelse og familiestruktur i barndommen. Vores resultater er i overensstemmelse med amerikanske studier og viser, at drenge har større gavn af et fordelagtigt familiemiljø end piger med hensyn til, hvordan de klarer sig i folkeskolen. Fædres uddannelse, som ikke berøres i tidligere studier, er særlig vigtig for sønner. Men vi finder et meget anderledes mønster af forældres indflydelse på arbejdsmarkedsudfald. Kønsforskelle i uddannelsesniveau, beskæftigelse og indtjening er stigende i mødres uddannelse til gavn for døtre. Fædres uddannelse mindsker kønsforskellene i uddannelsesniveauet (til gavn for sønner) og arbejdsmarkedsresultater (til gavn for døtre). Vi konkluderer, at forskelle i drenge og pigers adfærd i teenageårene er dårlige proxyer for forskelle i færdigheder, der er afgørende for langsigtede resultater.

Kapitel 4, “The Intergenerational Transmission of Time Preferences Persists Across Four Decades”, er udarbejdet i samarbejde med Thomas Epper. Vi studerer overførslen af tidspræferencer på tværs af generationer ved hjælp af et eksperimentelt valideret spørgsmål fra en spørgeskemaundersøgelse. Forældres og børns tidspræferencer måles fire årtier fra hinanden og fjerner dermed problemer med omvendt årsagssammenhæng. Vores resultater viser, at forældre overfører tidpræferencer til deres børn, og resultaterne ændres ikke, når vi korrigerer for et omfangsrigt sæt af baggrundvariable. Vi viser endvidere, at mødre differentielt og i højere grad overfører deres tidspræferencer til døtre i forhold til sønner, selv når vi sammenligner søskende fra den samme familie. Dette tyder på, at socialisering påvirker børns grad af utålmodighed. Styrken af forældrenes overførsel af tidspræferencer mindskes ikke, når børnene bliver ældre, hvilket understreger vedholdenheden i forældres påvirkningen af børns tidspræferencer.

Chapter 1 – Origins of Gender Norms: Sibling Gender Composition and Women's Choice of Occupation and Partner

Origins of Gender Norms:
Sibling Gender Composition and Women's Choice of Occupation and Partner *

Anne Ardila Brenøe[a,b]

April 27, 2018

Abstract

I examine how one central aspect of the childhood family environment—sibling gender composition—affects women's gender identity, measured through their choice of occupation and partner. Using Danish administrative data, I causally estimate the effect of having a second-born brother relative to a sister for first-born women. The results show that women with a brother acquire more traditional gender norms with negative consequences for their labor earnings. I provide evidence of increased gender-specialized parenting in families with mixed sex children, suggesting a stronger transmission of traditional gender norms. Finally, I find indications of persistent effects to the next generation of girls.

JEL classification: I2, J1, J3

Keywords: Gender identity, sibling gender, occupational choice, family formation.

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1 Introduction

Across most OECD countries, women today attain more education than men do and participate almost equally in the labor force (OECD, 2016; OECD, 2017). But why do women keep choosing fields of study leading to substantially lower-paid occupations (Blau and Kahn, 2016)? Although the barriers to women's participation in education and the labor force have been removed in the attempt to reach gender equality, gender identity still plays an important role for gender differences in behavior and subsequently in economic outcomes (Akerlof and Kranton, 2000; Bertrand, 2011; Goldin, 2014). To really understand why women continue behaving in ways leading to inferior labor market outcomes relative to the ones of men, we need to better understand the origins of—especially women's—gender norms. In this study, I focus on the importance of one key aspect of the childhood family environment—sibling gender composition—for women's socialization and development of gender conformity.

The family constitutes an essential facet of a child's socialization process. Parents act as important role models and transmit gender norms to their children (Farre and Vella, 2013; Fernández et al., 2004; Humlum et al., 2017; Johnston et al., 2013; Kleven et al., 2018). Siblings, at the same time, are close peers during childhood and often sustain long-lasting relationships throughout life (McHale et al., 2013). A child's birth order in the sibship influences, for instance, educational attainment and the development of personality traits through social family interactions (Brenøe and Molitor, 2018; Black et al., 2005, 2017; Lehmann et al., 2016). Sibling gender composition may additionally have a crucial impact on how siblings interact with each other as well as how parents interact with their children (McHale et al., 2003). Parents may, for instance, invest differently in their children depending on the children's gender composition which, in turn, could alter the intergenerational transmission of gender norms.

To examine how sibling gender composition affects the development of women's gender identity, I use high-quality administrative data for the total population in Denmark from 1980 through 2016. With this comprehensive data set, I evaluate women's gender identity through their revealed gender conformity in terms of their choice of occupation and partner from age 31 through 40 (proxied by the gender share in their own and their partner's occupations, respectively). To provide causal estimates of the impact of sibling gender, I exploit the random assignment of the second child's gender in families with a first-born daughter, conditional on the parents having a second child. The crux of my identification strategy is thus to compare the choices for first-born women with a second-born brother to those with a second-born sister. Sibling gender composition has a small impact on family size, yet, I show that family size is not a confounding factor for the effect of sibling gender composition on women's gender conformity. This empirical approach distinguishes itself from previous studies on sibling gender composition, as they generally include all siblings both in the measure of sibling gender composition and in the estimation sample.[1] Considering all siblings is problematic, however, as the final sibling gender composition in a sibship is endogenous. Therefore, studying the effects of older siblings' gender on younger siblings' outcomes may lead to selection bias. If parents, for example, decide to have a second child depending on their first child's gender and if parents with different gender preferences raise their children differently, the estimated effects would be biased. By focusing on the second-born child's gender, I avoid selection bias, as parents do not know the gender of their unborn child when deciding to have another child.

The setting for this study is ideal, as Denmark has been one of the front runners in terms of gender equality for decades. Women from the cohorts of study (1962–1975) attain slightly more education than men do[2] and importantly, labor force participation is not gendered. That labor market participation and family formation are not associated with gender identity is a unique (and very essential) feature for the empirical analysis, thereby removing concerns regarding selection into having an observation on choice of occupation, choice of partner, and the outcomes of a first-born child. Yet, pronounced gender differences in occupational choice still

[1] E.g. Amin (2009); Anelli and Peri (2014); Bauer and Gang (2001); Butcher and Case (1994); Conley (2000); Cools and Patacchini (2017); Cyron et al. (2017); Hauser and Kuo (1998); Kaestner (1997); Oguzoglu and Ozbeklik (2016); Rao and Chatterjee (2017). The only exceptions from such strategy are Cronqvist et al. (2015) and Peter et al. (2015), investigating the effect of a co-twin's gender on financial risk taking, education, earnings, and family formation. Moreover, Gielen et al. (2016) employ a difference-in-differences strategy to estimate the effect of having a male twin on earnings; yet, their interest is whether exposure to prenatal testosterone (rather than sibling gender composition per se) has an effect on earnings. Cools and Patacchini (2017) and Rao and Chatterjee (2017) both provide a robustness check of their estimates on wages in which they only consider the sex of a next younger sibling.

[2] This is a fortunate feature, as previous studies on sibling gender composition have been concerned with the potential role of differential parental monetary investment in daughters when also having sons, as parents in more traditional societies tend to favor boys. In Denmark, on average, parents do not favor one gender over the other (Andersson et al., 2006). Therefore, sibling gender is not associated with financial constraints.

persist. Women are, for example, still heavily underrepresented in occupations within Science, Technology, Engineering, and Mathematics (STEM). Therefore, the setting is, in many ways, comparable to the conditions faced by women in other developed countries today.

My results show that having a second-born brother relative to a sister increases first-born women's gender conformity: women with a brother work in more female-dominated occupations during their 30s and choose more traditional partners. In particular, women with a brother are 7.4 percent less likely to work within STEM. In other words, having a brother decreases women's probability of participating in traditionally male-dominated occupations. STEM is one important example of such occupations due to its potential consequences for the individual woman and society, given the higher wage returns to STEM fields and the need for a talented STEM workforce to sustain long-run economic growth (Altonji et al., 2015; Kirkebøen et al., 2016; Peri et al., 2015). Consistent with the fact that male-dominated occupations typically are better paid, I show that women with a brother earn less than those with a sister. I provide evidence that differences in labor market participation and family formation cannot explain the effects on occupational choice or labor earnings. While the main analysis concerns the development of women's gender identity, I also briefly present the results from a similar analysis for men (Section 6). Consistent with the findings for women, the results suggest that having an opposite sex sibling enhances men's gender identity.

The effect of sibling gender on women's gender conformity propagates through life and is already visible when considering their educational choice. While sibling gender has no effect on educational attainment or achievement, women with a brother complete less male-dominated educations. As an example, having a brother decreases women's probability of completing any field-specific STEM education by 11.3 percent. This effect on women's falling out of STEM fields is already present in their first educational choice after compulsory schooling at age 16. The key finding that women with a brother acquire more gender-typed human capital further motivates an analysis of whether the effects persist into the human capital formation of the next generation. Remarkably, the results show that daughters' comparative advantage in language over math in school is larger for those with a more gender-conforming mother, i.e. for daughters of mothers with a brother relative to daughters of mothers with a sister. Thus, I find striking evidence of very persistent long-run consequences of women's childhood family environment.

Why does sibling gender affect the development of women's gender identity? The effect of having a brother could go through either child-parent and/or child-sibling interactions.[3] I provide compelling evidence in favor of the former channel by showing that parents of mixed sex children invest their time more gender-specifically in their first-born daughter than parents of same sex children. The results from heterogeneity analyses further indicate that the effect of having a brother is largest for women from more traditional families. These findings are consistent with the argument, similar to the one put forward in the same sex education literature (Booth et al., 2013; Schneeweis and Zweimüller, 2012), that having an opposite sex sibling increases girls' exposure to gender-stereotypical behavior and thereby increases their inclination to acquire more traditional gender norms. In support of this argument, Cools and Patacchini (2017) and Rao and Chatterjee (2017) provide some indications that women with brothers hold more traditional gender attitudes than those without brothers.

My focus on the social environment and the origins of gender norms is consonant with recent studies that trace gender gaps in educational outcomes to factors such as teacher stereotypes, the gender of school peers and teachers, and parental and sibling role models.[4] A strand of the literature shows, for instance, that gender-stereotypes in the school environment affect the gender gap in math test scores.[5] Fewer studies, however, trace effects into outcomes with consequences for economic well-being in adulthood, such as field of education, working decisions, and earnings—in part, due to limited data availability. Some exceptions exists, however. For instance, Olivetti et al. (2016) show that having more female peers with working mothers during adolescence increases young women's probability of working and Kleven et al. (2018) show that women's child penalty on wages is largest for those from more traditional families. The literature on sibling gender composition is small and has predominantly been concerned with educational attainment, while

[3]The impact of having a brother on gender identity could also theoretically be due to changes in ability and parental resource constraints. However, I rule this out by showing that sibling gender does not affect school performance or attainment.

[4]See e.g. Anelli and Peri (2014, 2016); Bottia et al. (2015); Brenøe and Lundberg (2017); Brenøe and Zölitz (2018); Carrell et al. (2010); Cheng et al. (2017); Zölitz and Feld (2017); Humlum et al. (2017); Joensen and Nielsen (2017); Johnston et al. (2013); Oguzoglu and Ozbeklik (2016).

[5]Several studies find that having a gender-stereotypical teacher increases the math test score gap, mainly by decreasing girls' performance (Alan et al., 2017; Carlana, 2017; Lavy and Sand, 2015; Lavy and Megalokonomou, 2017).

a couple of more recent papers focus on wages.[6] The evidence on educational attainment is overall mixed, while studies on wages reach a more consistent finding that both male and female wages are negatively associated with having an opposite sex sibling—similar to my findings.

This paper makes five important contributions to the existing literature. First, I provide a comprehensive analysis of how sibling gender composition causally affects the development of women's gender identity, using two novel measures of gender conformity. Second, the large sample size and administratively reported occupations provide precisely estimated effects on the gender conformity of women's occupational choice.[7] Third, to the best of my knowledge, I am the first to consider the gender conformity of the choice of women's partner, which is again only possible due to the rich data set, as I am able to identify all partners and their occupations without relying on self-reports. Fourth, I document lasting effects to the next generation of girls, thereby stressing the persistence of gender norms. Fifth, I conduct a large quantitative analysis of how sibling gender composition affects child-parent interactions, thereby providing a detailed picture of an important channel through which the effects on gender identity operate.

2 Empirical Strategy

The aim is to estimate the causal effect of sibling gender composition on the formation of women's gender identity. Simply comparing women from families with different gender compositions would, however, not provide valid estimates of the causal effect of sibling gender composition due to selection. The final gender composition in a family is endogenous, as parents decide whether or not to have more children after each childbirth and thereby when knowing their current children's gender composition. If parents' decision to have a second child depends on the first child's gender and if such gender preferences also affect how parents raise their children, it is not possible to estimate the causal effect of "current" (first-

[6] See the references in Footnote 1. A general problem, though, is small sample sizes, often resulting in quite imprecise estimates, and potential biases.

[7] This is in contrast to the only few existing studies that have attempted to consider occupational outcomes, such as an occupational prestige score and binary indicators for occupational groups (Cools and Patacchini, 2017; Rao and Chatterjee, 2017). Their sample sizes ($< 5,000$) have, however, been too small to allow for any clear conclusions; the estimates of the signs are generally consistent with my main findings, though.

born) children's gender on "future" (second-born) children's outcomes because not all "future" children are born.[8]

To reach the goal of estimating the causal effect of sibling gender composition, I focus on the random assignment of the second-born child's gender. Because parents do not know the gender of a subsequent child when they make the decision to progress to the next parity, I *can* causally estimate the effect of a "future" child's gender on "current" children's outcomes. Thus, I leverage the random assignment of the second child's gender in families with a first-born daughter, conditional on having a second child. In other words, I compare first-born women who have a second-born brother to first-born women who have a second-born sister. Thereby, the identifying assumption is that conditional on the first child's gender and conditional on having a second child, the sex of the second child is random.

The empirical specification for the main analysis is:

$$Y_i^{First-Born} = \alpha_0 + \alpha_1 Brother_i^{Second-Born} + X_i' \delta + \nu_i, \qquad (1)$$

where $Y_i^{First-Born}$ measures woman i's (who is first-born) gender conformity. The estimate of interest is α_1, representing the effect of having a second-born brother. X_i is a vector of fixed effects for birth municipality, year-by-month of birth, spacing in months to the second-born sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education.[9] ν_i is the error term.[10]

As this strategy only relies on the random assignment of the second child's sex, parents can respond to the gender composition of their first two children in terms of subsequent fertility. Consistent with the literature exploiting sibling sex composition as an instrument for family size (e.g. Angrist and Evans (1998)), Appendix Table A1 shows that, for the main sample of the analysis (described in Section 3), having two mixed sex children reduces family size by 0.07 children, on average. Therefore, family size might mediate some of the effect of having a second-born brother if family size has an independent impact on gender identity. Existing studies find that family size does not affect educational attainment in Israel or Norway, using twins as an instrument for family size (Angrist et al., 2010; Black et al., 2005).

[8]Appendix A.1 shows the selection bias problem more formally and discusses other reasons for selection bias than parental gender preferences.

[9]If the parent does not have a field-specific education, I use their field of occupation.

[10]I do not cluster the standard errors; however, the results do not change if I do so.

In Appendix A.2.1, I replicate this finding in the Danish context and show that neither does family size affect the different measures of gender conformity. Appendix A.2.2 provides additional tests of the sensitivity of the findings, which further lend support to the conclusion that the results are robust to family size. Based on this wide battery of tests, family size does not seem to be an important confounder of the effect of sibling gender.

3 Data

3.1 Data and Sample Selection

I use Danish administrative data for the total population from 1980 through 2016. One central feature of this data set, compared to most previous studies on sibling gender composition, is that I can link all children to their parents and siblings. Thus, I observe parents' complete fertility history and thereby, correctly measure the sibling gender composition. Furthermore, I have information on parents' date of birth, length, type, and field of education, labor market attachment, and occupation. For the children, I annually observe labor market outcomes, educational enrollment and completion, fertility, cohabitation, and marital status. Finally, I observe the school performance of the children's children.

I restrict the sample to women born between 1962 and 1975 to be able to study the choice of occupation and partner when these women are in their 30s. Moreover, I only include first-born women, who are the first child to both the mother and father; I exclude immigrants;[11] I only consider individuals who have at least one full sibling (same mother and father) born less than four years apart and who survives the first year of life; I exclude families in which either the first or second child is a twin; and finally, I exclude those few women who die before age 40 or do not live in Denmark at any time between age 31 and 40 when the main outcome variables are measured.[12] I refer to this sample of first-born women as the *main sample*.

[11]For first-generation immigrants, I do not necessarily have complete sibling or parental information. Second-generation immigrants would have represented approximately one percent of the sample, reason for which I decided to exclude them to have a more homogeneous sample. However, including second-generation immigrants does not change the results.

[12]Sibling gender composition does not affect attrition due to these restrictions.

Table 1
Descriptive Statistics on Childhood Family Environment for Sample of First-Born Women

Panel A: Statistic by Gender of the Second-Born Sibling

	Sister		Brother		t-test
	Mean	SD	Mean	SD	p-value
	(1)	(2)	(3)	(4)	(5)
Predetermined Characteristics					
Spacing (months)	29.9	9.6	30.0	9.6	0.16
Mother's age at birth (years)	22.9	3.6	22.8	3.6	0.21
Father's age at birth (years)	25.7	4.4	25.6	4.4	0.06
Mother's education (years)	10.9	3.2	10.9	3.2	0.62
Father's education (years)	11.8	3.3	11.8	3.3	0.54
Mother has $\geq$ 12 years of education	50.8	50.0	51.2	50.0	0.28
Father has $\geq$ 12 years of education	65.7	47.5	65.8	47.4	0.85
Both parents have $\geq$ 12 years of edu	41.5	49.3	41.8	49.3	0.33
Mother in care or administration	15.6	36.3	15.8	36.4	0.42
Father in STEM	8.2	27.4	8.3	27.6	0.58
Mother in care/adm & Father in STEM	2.4	15.2	2.4	15.3	0.68
Parental Response to Sex Composition					
Number of siblings	1.7	0.9	1.6	0.9	<0.01
Has $\geq$ 2 siblings	39.9	49.0	34.6	47.6	<0.01
Has $\geq$ 3 siblings	8.4	27.8	7.1	25.6	<0.01
Lives with both bio parents	81.0	39.2	81.1	39.1	0.62
Lives with mother, sib with father	4.6	20.9	9.9	29.9	<0.01
Parents Equal Division of Labor	33.7	47.3	33.4	47.2	0.38
Observations	50,757		52,776		
Panel B: Balancing Test					
Joint F-statistic			0.92		
Prob > F			0.92		

Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Panel A shows the average and standard deviation of family background characteristics for first-born women with a second-born sister [Columns (1) and (2)] and brother [Columns (3) and (4)]. Column (5) reports the p-values from t-tests of significance between the averages of the two groups of women. All binary variables (variables measuring shares) are multiplied by 100 to express percent (percentage points). Panel B tests whether the control variables included in X_i in Equation (1) can predict having a second-born brother. F-test of joint significance of all control variables.

Table 1 provides descriptive statistics on the childhood family environment for the main sample by the gender of the second-born sibling. As expected, these women come from families with similar predetermined family characteristics regardless of sibling gender. On average, spacing to the younger sibling is 2.5 years, mothers are 22.9 years at birth and have 10.9 years of education, while fathers are 25.7 years and have 11.8 years of education. When it comes to characteristics that the parents can manipulate after realizing the gender composition of their first two children, we see that those with two daughters are more likely to have more children, as discussed in Section 2. Meanwhile, the probability of having both parents working equally[13] during childhood or living with both biological parents at age 17 does not differ by sibling gender composition. Among those not living with both parents at age 17, however, we see a clear difference in the family living arrangement: divorced parents with mixed sex children are more likely to live with their same sex child only.

To provide support for the identifying assumption, that sibling gender is random, Column (5) in Panel A tests whether the background characteristics differ by gender of the second-born sibling. Considering the predetermined characteristics, only father's age at birth differs marginally between the two groups.[14] Panel B shows statistics from a balancing test, testing whether the demographic characteristics included in X_i in equation (1) can predict sibling gender. More precisely, it reports the *F*-test of joint significance of all the covariates in a regression where the outcome is an indicator for having a second-born brother. The *F*-test strongly rejects joint significance. Thus, this balancing test supports the identifying assumption that the younger sibling's gender is random, conditional on the first child's gender and conditional on having a second child.[15]

[13] I define this as the tertile of families in which the parents' division of labor until the child turns 19 years is most equal. More precisely, fathers in this group work at most 62 percent of total parental labor supply. I observe parents' labor supply through a mandated pension scheme (ATP), in which employers contribute for each employee based on the number of hours worked.

[14] To account for this small baseline difference, I flexibly control for parental age among a wide range of other fixed effects in the analysis.

[15] The graphs in Appendix Figure A1 illustrate the estimates from an event study of the effect of having a second-born son on a variety of parental socio-economic characteristics. The gender composition of children does not affect parental cohabitation, marital status, length of education, employment, or annual labor earnings before or around the birth of their first child.

3.2 Outcome Variables

The three main outcome variables evaluate the degree of women's gender conformity. The first outcome reflects how gender-typed the individual woman's occupational choice is. More precisely, I construct this variable as the natural logarithm of the average male share in the woman's 4-digit occupation codes observed between age 31 and 40.[16] The second outcome measures the share of years between age 31 and 40 the woman works in a high-skilled STEM occupation. The third outcome quantifies how traditional the woman's choice of partner is. This variable measures the natural logarithm of the female share in the partner's occupation.[17] Table 2 provides descriptive statistics on the outcome variables for the main sample of women by sibling gender and for a sample of men, which is selected similarly to the main sample, for comparison. We observe a strong degree of gender segregation in occupational choice. While women, on average, have 33 percent men in their occupation, this number is 72 percent for men. Similarly, women's partners have, on average, 28 percent women in their occupation compared to 66 percent for men's partners. Moreover, men are three times more likely than women to work within STEM.

To study potential causes and consequences of occupational choice, I further consider educational and labor market outcomes. I examine labor market outcomes from age 18 through 40 in terms of the labor earnings percentile by age and cohort, work experience, and unemployment history. The earnings percentile provides a standardized measure of relative income that includes individuals with zero earnings, is comparable across cohorts and ages, and is constructed based on the total population. At age 40, women have an average earnings percentile of 49, corresponding to a mean labor income of 320,000 DKK (43,000 EUR). While women only earn 70 percent of men, men and women participate almost equally in the labor market: by age 40, women (men) have 14 (16) years of work experience and 1.8 (1.2) years of unemployment. Similarly, these cohorts of women and men attain

[16] I use the Danish version of International Standard Classification of Occupations (DISCO), which I observe from 1991 through 2013.

[17] I define the partner as the mode person with whom the woman cohabits or is married between age 31 and 41. Sibling gender has no impact on women's probability of having an observation on the partner's occupation (not reported). I consider the logarithm of the male share in the woman's own occupation and the logarithm of the female share in her partner's occupation because these measures best approximate a normal distribution rather than considering the logarithm of the male share in both persons' occupations.

Table 2
Descriptive Statistics on Outcome Variables for Sample of First-Born Women by Gender of Second-Born Sibling (and First-Born Men for Comparison)

	Women				Men	
	Sister		Brother		Sister/Brother	
	Mean (1)	SD (2)	Mean (3)	SD (4)	Mean (5)	SD (6)
Choice of Occupation and Partner						
Male share in own occupation	33.6	21.1	33.2	20.9	71.6	22.1
STEM occupation	5.2	19.0	4.8	18.1	14.2	30.3
Female share in partner's occ	28.4	21.4	28.0	21.4	66.4	20.3
Labor Market Outcomes at age 40						
Earnings Percentile	49.1	24.8	48.7	24.7	64.4	27.4
Earnings (1,000 2015-DKK)	320.6	197.6	318.6	197.8	460.7	395.3
Work experience (months)	168.9	63.4	168.6	63.7	192.2	69.0
Unemployment (months)	21.4	25.5	21.5	25.6	14.3	21.7
Education by age 30						
Male share in education	36.0	21.5	35.7	21.5	66.4	25.2
Length of education (months)	159.6	26.7	159.5	26.6	158.8	27.4
Academic high school GPA (std.)	0.02	0.99	0.01	0.99	0.09	1.03
Any STEM enrollment	8.3	27.6	7.7	26.7	41.6	49.3
Any STEM completion	5.1	21.9	4.5	20.8	30.3	45.9
Marital and Fertility History by age 41						
Cohabit share age 18–41	26.8	21.0	26.0	20.7	23.8	19.6
Married share age 18–41	39.0	27.6	38.9	27.7	30.1	25.5
Has any children	88.7	31.7	88.5	31.9	79.5	40.4
Number of Children	2.0	1.1	2.0	1.1	1.7	1.1
Age at first childbirth	27.3	4.7	27.3	4.7	29.3	4.6
First-Born Child's Grade 9 GPA (standardized with mean 0, SD 1)						
Daughter language	0.37	0.93	0.40	0.92	0.34	0.94
Daughter math	0.13	0.95	0.14	0.96	0.08	0.97
Son language	-0.07	0.96	-0.07	0.97	-0.11	0.97
Son math	0.23	0.95	0.24	0.95	0.19	0.97
Observations	50,757		53,012		108,366	

Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart); the sample of men corresponds to the one of women with the exact same sample selection criteria. Columns (1) and (3) show the average outcome variables for first-born women with a second-born sister and brother, respectively, while Column (5) shows the average for first-born men regardless of the second-born's gender. All binary variables (variables measuring shares) are multiplied by 100 to express percent (percentage points).

almost equal length of education; by age 30, women have on average completed 13.3 years of education and men have completed 13.2 years. Consistent with the differences in occupational choice, the male share in the highest completed degree is much lower for women (36 percent) than for men (66 percent) and women are much less likely to enroll in and complete any field-specific STEM education.[18]

Furthermore, I examine whether sibling gender affects family formation through age 41. This aspect of women's life might reflect a certain degree of gender-conformity and might at the same time influence labor market outcomes (Bertrand, 2011). First, I consider the share of years between age 18 and 41 during which the woman cohabits without being married (henceforth *cohabit*) and is married, respectively. Second, I consider the probability of having any children, the number of children, and age at first childbirth conditional on having any children. Although having a partner (and being married) and having children might reflect a greater degree of gender-stereotypical behavior, it is not inevitably the case (Bertrand et al., 2016). Cohabitation could instead reflect non-traditional behavior, as marriage is the tradition. Moreover, the vast majority (89 percent) of women have at least one child and most of those having children have exactly two. Therefore, gender identity may not necessarily influence family formation.

Finally, the last group of outcomes concerns the school performance of the next generation. For this, I consider the outcomes of the first-born child and split the sample by the child's gender.[19] I examine the externally-graded grade point average (GPA) from the grade 9 written language (Danish) and math exams. Both measures are standardized with mean zero and standard deviation (SD) one by exam year for the entire student population. Generally, and as seen from the data, girls perform much better (0.45 SD) than boys in languages, while boys perform slightly better (0.10 SD) than girls in math. Therefore, languages may be perceived as more feminine and math more masculine. Thus, if mothers' gender identity transmits to their children (daughters), we might observe a widening in the gap between language and math performance. Given previous findings, suggesting that mothers influence their daughters more than sons and vice versa for fathers

[18]See Appendix A.3 for details on the educational outcomes and the educational system in Denmark with emphasis on STEM education.

[19]Given child gender is independent of the gender of the mother's sibling, this split does not create any bias. Yet, sibling gender might affect the mother's gender preference for her own children and thereby her subsequent fertility choices. Therefore, I only consider women's first-born children. I do not observe any selection into having an observation on a first-born child's outcomes.

(Brenøe and Lundberg, 2017; Brenøe and Epper, 2018; Humlum et al., 2017), we would mainly expect to observe an effect of the gender of the mother's sibling on daughters' and not on sons' performance.

4 Results

4.1 Gender Identity: Choice of Occupation and Partner

Table 3 shows the main results on the impact of sibling gender on women's choice of occupation and partner, with different control versions. The models in Column (1) show the raw means between first-born women with a second-born sister and those with a second-born brother, while Column (2) includes basic demographic controls. Column (3), the preferred model, further controls for parental education. Finally, Column (4) flexibly adds controls for family size and the sex of potential third- and fourth-born siblings.[20] As family size is an outcome of sibling gender composition, the latter control version might bias the estimates. This control version, however, works as a robustness check of the results, as family size might also be considered a confounding variable. Regardless of the covariates included, the estimates across the different control versions are almost identical, supporting the assumption that sibling gender is random and illustrating that family size is not a principal mediator of the effect of sibling gender (as discussed in more detail in Appendix A.2). Therefore, the rest of this paper proceeds by presenting the results, using the preferred control version in Column (3).

Overall, the results show that having a second-born brother relative to a sister enhances women's gender identity. First-born women with a second-born brother work in occupations with 1.22 percent fewer men compared to first-born women with a second-born sister. Note that this difference in occupational choice is observed well into these women's labor market careers during their thirties (as an average from age 31 through age 40). Consistent with this, having a brother also reduces women's probability of working within STEM by 0.38 percentage points, corresponding to a decrease of 7.35 percent relative to the mean for women with a sister. Consequently, the results clearly show that having a brother induces women to exhibit more traditional choices of occupation. In other words, they are less

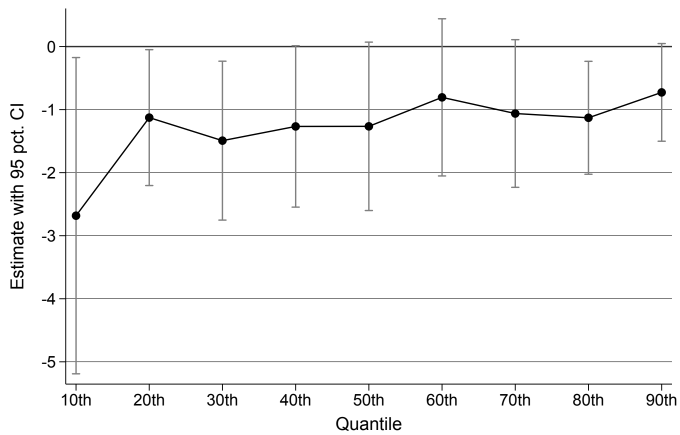
[20]The estimates are identical when not controlling for third- and fourth-born siblings' gender.

Table 3
Effect of Sibling Gender on Choice of Occupation and Partner

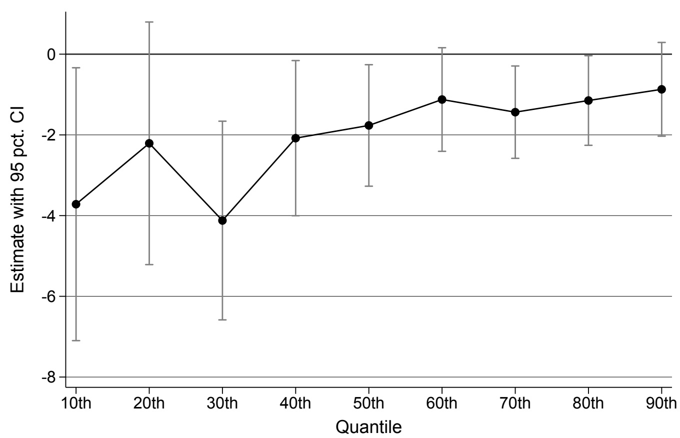
	(1)	(2)	(3)	(4)
Panel A: Log(Male Share in Own Occupation)				
Second-Born	-1.16**	-1.17**	-1.22***	-1.29***
Brother	(0.48)	(0.47)	(0.47)	(0.47)
Observations	103,769	103,769	103,769	103,769
Panel B: Share of Years Working in STEM Occupation				
Second-Born	-0.38***	-0.37***	-0.38***	-0.42***
Brother	(0.12)	(0.11)	(0.11)	(0.11)
Observations	103,769	103,769	103,769	103,769
Panel C: Log(Female Share in Partner's Occupation)				
Second-Born	-1.98***	-1.74***	-1.88***	-1.89***
Brother	(0.67)	(0.67)	(0.66)	(0.66)
Observations	95,087	95,087	95,087	95,087
No controls	✓			
Basic controls		✓	✓	✓
Parental education			✓	✓
Family size				✓

All estimates are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. *Basic controls* include fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, and paternal age at birth. For the own occupation outcomes, basic controls also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. For partner's occupation, basic controls also include dummies for the partner's number of occupational observations and age at first and last observation. *Parental education* controls include fixed effects for maternal level-by-field of education and paternal level-by-field of education. *Family size* controls include dummies for the number of biological siblings and dummies for the number of children the mother and father potentially have, respectively, from later relationships, and the gender of potential third- and fourth-born siblings. The occupational outcomes of the first-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41.

Figure 1
Distributional Effects of Sibling Gender on Choice of Occupation and Partner



(a) Log(Male Share in Own Occupation)



(b) Log(Female Share in Partner's Occupation)

All estimates are multiplied by 100 to express effects in log-points. The whiskers represent the 95 percent confidence interval. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). All estimates come from separate quantile regressions. All models control for quadratic spacing to the second-born sibling, mother's and father's cubed age at birth, and absorb fixed effects for year of birth, indicators for missing parental age information, and a constant. The models in Graph (a) further control for dummies indicating the number of occupational observations and the models in Graph (b) control for the partner's number of occupational observations and age at first and last observation.

prone to opt into traditionally male-dominated occupations, of which STEM is one relevant example.

Moreover, sibling gender has a significant impact on the choice of partner in terms of the degree of how gender-typed his occupation is. Having a brother rather than a sister induces women to choose a partner who works in more male-dominated occupations. On average, women with a brother have a partner working in occupations with 1.88 percent fewer women than women with a sister. Not reported, having a brother increases the difference in the male share between the woman's own and her partner's occupations by 0.80 percentage points. These results hereby demonstrate a powerful effect of having a brother on women's choice of gender-stereotypical occupations and partners.

Figure 1 considers whether the effects differ across the different parts of the distribution, by presenting the results from quantile regressions. Both for the male share in the woman's own occupation and the female share in the partner's occupation, the estimates are not statistically significantly different from each other from the tenth through the ninetieth percentiles. Yet, for both measures, the estimates indicate largest effects at the lower part of the distributions. At the tenth percentile, the estimated effects of having a brother relative to a sister are approximately twice the magnitude of the ones seen in Table 3. This suggests that those women who are affected the most by having a brother are those who are more traditional than the average.

If the effect of sibling gender, at least partly, goes through the way in which parents treat their children, we might observe some heterogeneity in the effect of having a brother by parental characteristics.[21] Panel A in Table 4 includes an interaction term between sibling gender and an indicator for having parents working (close to) equally during childhood. Remarkably, the effect of having a brother on occupational choice disappears for women coming from more gender-equal families. This suggests that women with more gender-stereotypical parents drive the effect of sibling gender on choosing more female-dominated occupations. Moreover, the results in Panel B suggest that the effect of having a brother is largest for those women with more traditional parents in terms of their educational field. The effects seem to be largest in magnitude for those with a mother who has an academic education within care or administration and for those with a father who has an academic education within STEM.

The effect of having a brother is, furthermore, the largest for those with at least

[21] As seen in Table 1, these parental characteristics do not differ by sibling gender composition.

Table 4
Heterogeneity: Choice of Occupation and Partner

	Log(Male Share in Own Occupation) (1)	Share of Years in STEM Occupation (2)	Log(Female Share in Partner's Occupation) (3)
Panel A: Parental Division of Labor During Childhood			
Second-Born	-1.63***	-0.36**	-1.96**
Brother (SBB)	(0.59)	(0.14)	(0.82)
SBB $\times$Equal	1.74*	-0.03	0.09
	(1.01)	(0.25)	(1.42)
Observations	100,020	100,020	91,706
Panel B: Parental Field of Academic Education			
Second-Born	-0.64	-0.19	-1.74**
Brother	(0.54)	(0.13)	(0.75)
SBB$\times$Mother Care/Adm	-1.46	-1.02***	-0.90
	(1.41)	(0.35)	(1.99)
SBB$\times$Father STEM	-3.79*	-0.75	-1.19
	(2.04)	(0.50)	(2.87)
SBB$\times$Mother Care/Adm	2.01	1.09	1.60
$\times$Father STEM	(3.91)	(0.96)	(5.54)
Observations	100,772	100,772	92,406
Panel C: Parental Years of Education			
Second-Born	0.84	-0.21	-1.64
Brother	(0.96)	(0.24)	(1.35)
SBB$\times$Mother≥ 12&Father< 12	-3.06*	-0.35	4.05
	(1.84)	(0.45)	(2.58)
SBB$\times$Mother< 12&Father≥ 12	-2.95**	-0.07	0.64
	(1.36)	(0.33)	(1.91)
SBB$\times$Mother≥ 12&Father≥ 12	-2.32*	-0.29	-1.99
	(1.21)	(0.30)	(1.70)
Observations	100,772	100,772	92,406

All estimates are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. Columns (1) and (2) also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. Column (3) also includes dummies for the partner's number of occupational observations and age at first and last observation. The occupational outcomes of the first-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41.

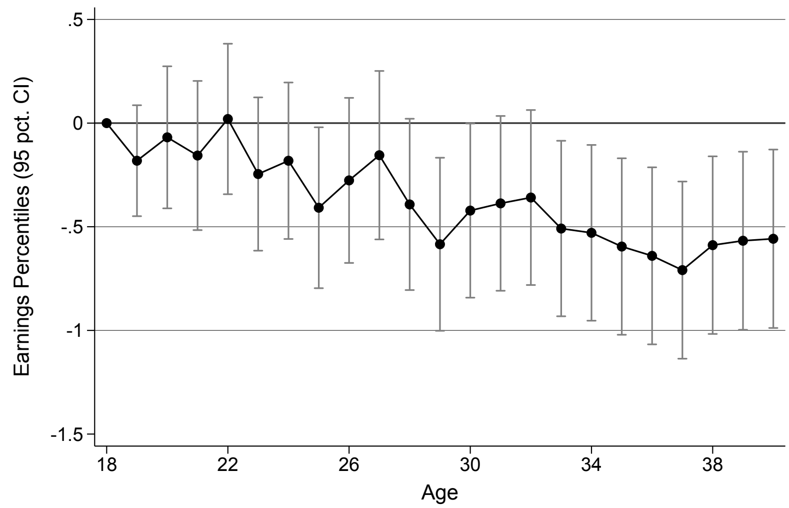
one highly educated parent ($\geq$ 12 years of education) for occupational choice. A highly educated parent will in most cases also imply having a parent with human capital that is traditionally associated with his or her own gender. For instance, most mothers with long education are within care and administration (e.g. nurse, secretary, and office work) and most fathers are within STEM. Therefore, these results again support the previous findings that the effect of having a brother is largest for those with more gender-stereotypical parents. Notably, the results also show that women with both parents having less education do not experience an effect of sibling gender. This suggests that the effect is not due to resource constraints, which has been put forward as a relevant mechanism in the sibling gender composition literature on educational attainment (Amin, 2009; Butcher and Case, 1994). Although the estimates are more imprecisely estimated for the other two outcomes, they are qualitatively consistent with the findings for the male share in the woman's occupation.

Expanding the sample to include women with up to eight years to their second-born sibling shows that sibling gender does not have an impact for those with long spacing to their sibling [Appendix Figure A2]. Though, the estimated effects by spacing are not statistically significantly different from each other, probably due to the small fraction of children with long spacing to their second-born sibling. This finding that individuals with long spacing to their younger sibling do not experience an effect of sibling gender might indicate the importance of sibling interactions. However, it could also be because parents with children spaced far apart treat the first-born child similarly regardless of the younger sibling's gender.

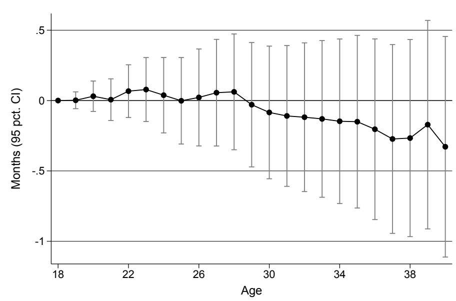
In sum, these heterogeneities indicate that the effect of having a brother is largest for women from more traditional families. This, in turn, suggests that differences in child-parent interactions are important for the effects of sibling gender composition on the formation of women's gender identity. Ceteris paribus, we would expect that parents with more gender-stereotypical human capital would reinforce gender-specialization to a larger extent than those parents with less gender-specific human capital (Humlum et al., 2017). Additionally, we would expect that spending more time with the mother than with the father would influence the child more in the direction of the mother's (female) than the father's (male) interests. Therefore, the results are consistent with the hypothesis that parents of mixed sex children invest more time in their same sex child than parents of same sex children; Section 5 elaborates more throughly on this.

4.2 Labor Market Outcomes

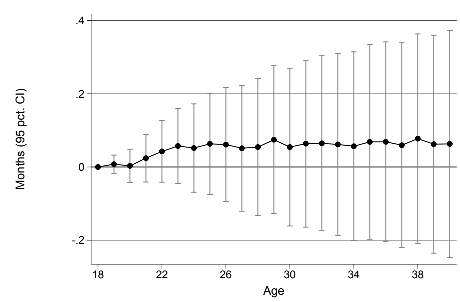
Figure 2
Effect of Sibling Gender on Labor Market Outcomes Age 18–40



(a) Earnings Percentile



(b) Work Experience



(c) Unemployment

Main sample (first-born women born 1962–1975 with a younger biological sibling born within four years apart). The whiskers represent the 95 percent confidence interval. All graphs illustrate the estimates from an event study of the effect of having a second-born brother, where age 18 forms the base. All models absorb time-specific fixed effects and individual fixed effects. *Earnings Percentile* measures the labor earnings percentile by age and cohort. *Work Experience* measures the cumulated lifetime work experience in months. *Unemployment* measures the cumulated lifetime unemployment in months.

As female-dominated occupations typically pay lower wages, an important consequence of the results on occupational choice may be reflected in lower labor earnings. To study this, I conduct an event study of the effect of having a brother on women's earnings percentile from age 18 through age 40 with age 18 as the base, controlling for individual fixed effects. Note, that sibling gender has a tight zero impact on earnings at age 18 (not reported). Once women enter the labor market,[22] we observe a negative effect of having a brother on the earnings percentile in the order of 0.5 percentile [Graph (a), Figure 2].[23] Such negative impact on earnings might be driven by differences in labor market participation rather than, or in addition to, occupational choice. This is, nevertheless, not the case, as Graphs (b) and (c) in Figure 2 illustrate that sibling gender does not affect women's cumulated work experience or cumulated length of unemployment. These findings of no effect on labor market participation (in terms of hours worked and unemployment) stress that labor market participation is not gendered in Denmark.

The finding of a negative consequence for earnings is not surprising, given the previous results of a lower participation in more male-dominated and STEM occupations. Similarly, Cools and Patacchini (2017) show that women in the U.S. with any brother earn less around age 30. Rao and Chatterjee (2017) do not find a significant effect of sibling gender composition on women's earnings among slightly older cohorts in the U.S., although their estimate of the effect of having a next younger male sibling indicates a negative impact. In contrast, studying a sample of female twins born in the first half of the last century, Peter et al. (2015) do not find an impact of having a co-twin brother on earnings. Moreover, both Cools and Patacchini (2017) and Rao and Chatterjee (2017) do not find significant effects of sibling gender composition on the type of occupation. This might be due to some important empirical limitations (as the sign of their estimates broadly support my findings) because these studies rely on much smaller sample sizes, self-reported measures of occupation, and their methodological approach (i.e. the inclusion of all siblings in the measure of sibling gender composition and the inclusion of all birth orders in the sample).

[22]Brenøe and Lundberg (2017) show, using Danish data, that almost everybody will have finished their education around age 30.

[23]Appendix Figure A3 demonstrates that the picture is similar when instead considering the earnings level and the natural logarithm of earnings.

4.3 Education and Family Formation

Another reason for the lower earnings could be due to differences in the accumulation of human capital. I do not find any evidence of an impact of sibling gender on educational attainment or school performance [Columns (2) and (3) in Panel A, Table 5].[24] Likewise, Cyron et al. (2017) does not find an effect of sibling gender on girls' cognitive or non-cognitive skills in first grade in the U.S.[25] Thus, sibling gender does not seem to affect differences in ability or (financial constraints in terms of) access to education. Consequently, these results demonstrate that sibling gender composition does not affect educational achievement or attainment, supporting an interpretation that changes in interests or identity are the channels of the effects of sibling gender on occupational choice. In contrast, the only existing study with causal estimates of sibling gender on educational attainment finds that having a male co-twin increases women's length of education (Peter et al., 2015). However, their sample might not be comparable to the more general population of singletons and for later birth cohorts.

While sibling gender does not affect overall educational attainment, the effect of sibling gender on occupational choice is closely mirrored in field of education by age 30. Having a brother reduces the share of men in the highest completed field-by-level of education by 1.36 percent.[26] Similarly, women with a brother relative to those with a sister are respectively 7.6 and 11.3 percent less likely to ever enroll in and complete any field-specific STEM education. Appendix Table A5 further shows that the effect is present already in the type of first educational enrollment after compulsory education and that it is seen for STEM degree completion at different levels of education. Thus, having a brother pushes women out of traditionally male-dominated fields as early as age 16 and is both seen in field of education as

[24] Not reported, sibling gender does not affect the probability of having an observation on high school GPA or the probability of enrolling or completing different levels of education. Appendix Table A5 further shows that there is no effect on different types of ability, measured through grade 9 language and math written exam GPA. Appendix Figure A4 further illustrates the distributions of the three GPA measures by sibling gender composition. The differences by sibling gender are extremely small; thus, distributional effects do not seem to be important.

[25] Similarly, I do not find any effect of sibling gender on personality traits [Big Five, growth mindset, trust, hedonism] or mental health [Strength and Difficulties Questionnaire (SDQ)] (not reported), based on the DALSC sample introduced in Section 5.

[26] Despite large changes in society over time, the effect of sibling gender on the male share in the highest completed education by age 30 does not differ systematically by decade of birth when including cohorts born though 1986 (not reported). This is consistent with the finding by Haines et al. (2016) that gender-stereotypes have not changed over the last three decades in the U.S.

Table 5
Effect of Sibling Gender on Education and Family Formation

	(1)	(2)	(3)	(4)	(5)
Panel A: Education by age 30					
	Log(Male Share)	Length (months)	High School GPA	STEM Enrollment	STEM Completion
Second-Born Brother	-1.36***	-0.12	-0.01	-0.63***	-0.57***
	(0.53)	(0.15)	(0.01)	(0.17)	(0.13)
Observations	103,541	103,562	47,588	103,769	103,769
Panel B: Family Formation by age 41					
	Cohabit 18–41	Married 18–41	Has Any Children	# of Children	Age at First Birth
Second-Born Brother	-0.80***	-0.12	-0.20	0.00	0.07**
	(0.13)	(0.17)	(0.20)	(0.01)	(0.03)
Observations	103,769	103,769	103,769	103,769	91,953

Estimates in Columns (1), (4), and (5) in Panel A and Columns (1), (2), and (3) in Panel B are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. The educational outcome models, except for high school GPA, further control for age at last observation in the education registry. *Log(Male Share)* measures the natural logarithm of the share of men in the highest completed education (narrow field-by-level) by age 30. *Length* measures the length of the highest completed education in months by age 30. *High School GPA* measures final GPA from the academic high school and is standardized by track and year of graduation for the total population with mean zero and standard deviation of one. *STEM Enrollment* indicates whether the woman has ever enrolled in a field-specific STEM education at age 16–27. *STEM Completion* indicates whether the woman has ever completed a field-specific STEM education by age 30. *Cohabit* measures the share of years age 18–41 during which the woman has cohabited with a partner without being married. *Married* measures the share of years age 18–41 during which the woman has been married. *Has Any Children* indicates whether the woman has at least one child by age 41. *# of Children* measures the number of children the woman has by age 41. *Age at First Childbirth* measures the age at the woman's first childbirth in years, conditional on having any children.

well as occupation.

The magnitude of the effects are comparable to previous studies examining the impact of various aspects of the social environment in school on study choice (Bottia et al., 2015; Carrell et al., 2010; Schneeweis and Zweimüller, 2012; Fischer, 2017). Moreover, the results are broadly comparable to other studies examining correlations between sibling gender composition and field of college major (Anelli and Peri, 2014; Oguzoglu and Ozbeklik, 2016). Appendix Table A10 displays the associations between gender of a first-born sibling and second-born women's gender identity, indicating similar but less robust correlations compared to the main results. These results are also closer to the ones in Anelli and Peri (2014) who do not find a significant association for women's enrollment in high-earnings college majors (although the magnitude of their estimate is relatively large). This stresses the importance of rigorously considering selection bias when the aim is to evaluate the causal effect of sibling gender.

In addition to differences in occupational and educational choice, one potential explanation for the negative effects on earnings might be differences in family formation. On one hand, due to the acquisition of more traditional gender norms, one might expect women with a brother to marry earlier, have children earlier, and have more children than women with a sister. However, such a conjecture implicitly requires that being married and having children is an important aspect of women's gender identity. This might very well not be the case in a modern setting in which women do not face a conclusive choice between having a family and a career (Bertrand et al., 2016; Goldin and Katz, 2002). The cohorts of women under study have, for instance, all had access to contraceptives, abortion, various family leave policies, and infant child-care options.[27] On the other hand, women with a younger sister might experience more competition in terms of being the first among the two who marries and has children, as men on average are older when they start their family formation. These two opposing forces might explain why I essentially do not find any effect of sibling gender on various aspects of family formation [Panel B in Table 5], consistent with the findings in Peter et al. (2015).

In terms of family formation, the results only suggest a small effect on cohabitation. Women with a brother cohabit 3.0 percent fewer years than those with a sister between age 18 and 41. This could, in fact, be due to more traditional

[27]Oral contraceptives (the pill) have been on the Danish market since 1966. All women have had free access to abortion since 1973 in Denmark.

gender norms, as more traditional women might want to wait longer before moving together with a partner before marriage.[28] Sibling gender has no effect on the probability of being married [Column (2)], age of first marriage, the probability of divorce, or age at first divorce (not reported). Thus, the only difference between women with a brother and those with a sister is that the former move together with a partner before marriage slightly later. This might explain the small positive (though negligible) effect on age at first childbirth. Overall, sibling gender has no effect on the fertility rate through age 41, i.e. close to complete realized fertility. Therefore, the effects of sibling gender on family formation are not a likely mediator of the effects on earnings. This, in turn, supports an interpretation of a causal positive effect of having a more male-dominated education and working in more male-dominated occupations on female earnings.

4.4 Persistent Effects to the Next Generation (of Girls)

Table 6
Effect of Sibling Gender on First-born Children's Grade 9 Performance

	Daughters		Sons	
	Language (1)	Math (2)	Language (3)	Math (4)
Second-Born Borther	2.37** (1.05)	0.19 (1.09)	0.36 (1.10)	0.33 (1.09)
Observations	29,047	29,036	29,262	29,262
Average	39.3	13.1	-6.0	23.8

All estimates are multiplied by 100 to express effects in percent of a standard deviation. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. First-born children to the main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart) born 1986–1999. All models absorb fixed effects for the mother's birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. The Grade 9 GPA measures come from the written exam at the end of grade 9 in respectively Danish and Math and are standardized by year of graduation for the total population with mean zero and standard deviation of one.

So far, I have documented that the childhood family environment affects the de-

[28]The majority of these cohorts cohabit and have children before marriage. Ninety-one percent of the women in the sample have cohabited at least one year before the year they get married and 53 percent get married in the year of their first childbirth or later.

velopment of women's gender conformity. Having a brother influences the family environment to such a degree that women choose more female-dominated occupations and more gender-conforming partners. An intriguing question is whether this effect on gender identity is sufficiently strong to affect the next generation—and in particular, the next generation of girls. To investigate this, I examine the school performance of these women's first-born daughters and sons, separately. If having a more gender-stereotypical mother (and father) affects the next generation, we would expect daughters to perform better in languages and/or worse in math. For, boys the prediction is less clear, as the literature typically finds boys to be less sensitive to the social environment (Bottia et al., 2015; Carrell et al., 2010; Fischer, 2017). Remarkably, Table 6 shows that daughters whose mother's second-born sibling is male relative to female perform 2.37 percent of a standard deviation better in languages, while there is no effect on their math performance or for sons. Thus, daughters' difference in language and math ability is larger for those with a more gender-conforming mother. This increase in girls' absolute advantage in languages over math might, in turn, predict more traditional choices of field of education. Notably, I find evidence of very persistent long-run consequences of women's childhood family environment.

5 Gender-Specific Parenting as a Relevant Mechanism

5.1 Literature Background

The previous section documents that sibling gender does matter for women's acquisition of traditional gender norms and that the effects seem to be largest among women from more gender-stereotypical families. This subsection draws on the literature to identify relevant mechanisms behind these findings, while the subsequent subsection provides some empirical evidence. Overall, I consider changes in identity to be the main channel of the impacts on choice of occupation and parter, as the previous analysis does not suggest that differences in educational attainment, ability, labor force participation, family size, or resource constraints are important or driving mechanisms. Consistent with the same sex education literature (Booth et al., 2014; Schneeweis and Zweimüller, 2012), the overarching argument is that girls with a brother are more exposed to gender-stereotypical behavior in the family and are therefore more inclined to acquire traditional gender norms. In this context, gender-stereotypical behavior could become more salient through changes in the nature of either child-sibling and/or child-parent interactions, including parental investments.[29]

First, parents might interact differently with their children depending on the gender composition in terms of quantity, quality, and content of time spent together. Assuming that both parents spend at least some time with their children, a traditional household specialization model suggests that parents gender-specialize their investment in children when having mixed sex children if mothers are more productive in creating female human capital and fathers are more effective in creating male human capital (Becker, 1973). Parents might also derive more utility from spending time with a same compared to an opposite sex child due to the type of activities done with the child. In both cases, parents of mixed sex children would gender-specialize, to a greater extent, than parents of same sex children.

McHale et al. (2003) suggest that because parents of mixed sex children have the opportunity to gender-differentiate their parenting, children with opposite gender siblings might have the strongest explicit gender-stereotypes. Endendijk et al. (2013) find some evidence that fathers with mixed sex children exhibit stronger gender-stereotypical attitudes than fathers with same sex children. Previous research has further documented that, overall, mothers talk more in general and more about interests and attitudes with daughters than sons (Maccoby, 1990; Leaper et al., 1998; Noller and Callan, 1990). Fathers, in contrast, talk more and spend more time with sons than daughters and have a greater emotional attachment to sons (Bonke and Esping-Andersen, 2009; Morgan et al., 1988; Noller and Callan, 1990). These different pieces of evidence thus suggest that parents of mixed sex children gender-specialize their parenting more and thereby expose their children more to gender-stereotypical behavior than parents of same sex children, which in turn might result in a stronger transmission of gender norms in families with mixed sex children.

Second, first-born girls might interact differently with their second-born sibling depending on the siblings' gender. In particular, having a brother might make girls more aware of "appropriate" female behavior and thereby induce them to de-

[29] Appendix A.4 provides a short overview of alternative mechanisms discussed in previous papers on sibling gender composition. These mechanisms cannot be the dominating explanations, as they are not compatible with the empirical findings.

velop more gender-stereotypical attitudes. For instance, Booth and Nolen (2012) show that girls attending same sex schools are no more risk averse than boys, while girls attending mixed sex schools are significantly more risk averse. Women are generally less competitive than men and this gender difference in competitiveness seems to be larger in mixed sex relative to same sex environments (Bertrand, 2011; Niederle and Vesterlund, 2011). Traditionally male-dominated (STEM) fields are further considered more competitive (Buser et al., 2014). Therefore, having a brother instead of a sister might change women's degree of competitiveness and thereby their preferences for working in competitive environments. Having a brother might thereby induce women to develop more gender-stereotypical attitudes due to a greater awareness of gender through sibling interactions. This, in turn, could be reinforced by parents' increased gender-specialization. In particular, previous studies have documented that women with brothers behave more family-centered and express more traditional attitudes towards gender roles (Cools and Patacchini, 2017; Rao and Chatterjee, 2017).

Thus, a particularly important mechanism for the observed effect of sibling gender on women's formation of gender identity—that I am able to test for empirically—is differences in child-parent interactions and, in particular, increased gender-specialization in families with mixed sex children. In the remainder of this section, I explore this mechanism by investigating the impact of sibling gender composition on parental time investment. More precisely, in the daily child-parent interactions, we might observe that parents of mixed sex children invest more quality time in their same sex child. This could explain the heterogeneity in the effect of sibling gender documented in Table 4. Furthermore, in the case of parental divorce, we might expect that children from mixed sex child-families would be more likely to live with their same sex parent compared to same sex children due to a larger degree of gender-specialized parenting. Consequently, common for these predictions is that a parent of mixed sex children influences his or her same sex child more than a parent of same sex children.

5.2 Empirical Evidence on Gender-Specific Parenting

To investigate whether sibling gender composition affects child-parent interactions, I draw on the Danish Longitudinal Survey of Children (DALSC).[30] The survey consists of five waves of children born in 1995 and is unique due to its very detailed information on parental time use and family socio-economic characteristics. For this analysis, I select first-born girls who have a second-born sibling born within four calendar years apart.[31] At age 7 and 11, both parents report how often they do different types of activities together with their first-born daughter. I construct an index on parental time investment, using principal component analysis, and standardize it with mean zero and standard deviation of one [Appendix Table A6]. I define quality time as playing with the child, helping with homework, doing out-of-school activities, reading/singing, and going on an excursion.

Columns (1) through (4) in Table 7 provide the results on parental time investment by each parent for the two ages, separately. Mothers of a first-born daughter and a second-born son invest *more* time in their first-born daughter at both ages compared to mothers with two daughters. The increase is in the magnitude of 14–17 percent of a standard deviation. In contrast, fathers invest 20–23 percent of a standard deviation *less* time in their first-born daughter when having mixed sex children. This reduction in total paternal time investment is driven by decreased time spent helping with homework and reading for the daughter [Appendix Table A7]. This finding indicates that girls with a younger brother receive less qualified help with homework in traditionally male-dominated subjects, which might prevent them from growing interests in these fields. This effect on father-daughter interactions furthermore translates into a substantially worse relationship between fathers and their first-born daughters when the second-born child is male relative to female [Appendix Table A8]. Overall, girls receive the same amount of time investment regardless of their younger sibling's gender. These results clearly show that first-born girls with a second-born brother experience more gendered parenting relative to those with a younger sister.[32]

[30]The study was designed by researchers from SFI, the Danish National Centre for Social Research, in collaboration with other research institutions. The survey consists of 6,011 randomly sampled children born between September and October, 1995 to a mother with Danish citizenship and consists of five waves (1996, 1999, 2003, 2007, and 2011).

[31]I only observe the year of birth of siblings and do therefore not have more precise information on spacing.

[32]For first-born boys, the overall picture is similar (not reported). Note, I cannot distinguish between

Table 7
Effect of Sibling Gender on Parental Time Investment in First-Born Daughters and Family Structure

	Parental Time Investment (Born 1995)				Family Structure (Born 1962–75)	
	Mother		Father		Lives w Both Parents	Lives w Mother & Sib w Father
	Age 7	Age 11	Age 7	Age 11		
	(1)	(2)	(3)	(4)	(5)	(6)
Second-Born	0.14*	0.17**	-0.20**	-0.23**	0.11	5.30***
Brother	(0.08)	(0.09)	(0.10)	(0.10)	(0.23)	(0.38)
Observations	594	562	421	415	102,137	19,196
Average	-0.00	0.00	-0.00	0.00	81.1	7.3
DALSC Sample	✓	✓	✓	✓		
Main Sample						
All					✓	
Divorced						✓

Standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. DALSC sample: Columns (1) through (4). Main sample: Columns (5) and (6). Each Column represents the results from separate regressions. All models using the DALSC sample control for (quadratic) mother and father's age and fixed effects for spacing to the younger sibling in years, parental marital status in 1996, parents having been together for at least 5 years in 1996, region of birth, maternal level of education, paternal level of education, and family income level in 1995. Both models using the main sample absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, paternal level-by-field of education, and age at observation of family structure. *Parental time investment* is constructed, using principal component analysis based on reports on how often each parent does certain quality time activities (playing, doing homework, doing out-of-school activities, reading/singing, going on an excursion) together with the child at a weekly basis and is standardized with mean zero and standard deviation of one; see Appendix Table A6. *Main Sample All* includes everybody who lives with at least one biological parent, while *Main Sample Divorced* excludes those living with both biological parents. *Lives w Both Parents* indicates that the first-born daughter lives with both biological parents at age 17. *Lives w Mother & Sib w Father* indicates that the first-born daughter lives with her mother and the second-born child lives with the father at age 17.

Ideally, I would have had similarly detailed data on parental inputs for the main sample. Such information is, however, not observed in the administrative registries. Instead, I do observe all children's family structure at age 17.[33] Sibling gender composition does not alter the probability of living with both biological parents [Column (5) in Table 7]. In the case of parental divorce or separation (henceforth *divorce*), the living arrangement between parents and children in the main sample might additionally shed light on child-parent interactions in terms of splitting parents' time. If parents of mixed sex children gender-specialize more than parents of same sex children, we would expect that divorced families with mixed sex children would be more likely than families with same sex children to have a living arrangement in which the first-born daughter lives with her mother and the second-born child lives with the father.

Conditional on living in a divorced family, the results show a pattern consistent with the prediction [Column (6)]. First-born daughters with a second-born brother are 5.30 percentage points (115 percent) more likely to live with their mother while their younger sibling lives with the father. These results consequently show a strong effect on the living arrangement among non-traditional families, thereby lending support to the previous findings (based on the much smaller DALSC sample) on more gender-specific parenting and time investment in families with mixed sex children. In conclusion, these findings support the hypothesis that parents of mixed sex children gender-specialize their parenting more than parents of same sex children, thereby strengthening the transmission of traditional gender-specific interests.[34]

whether this increase in gender-specialization is driven by changes in demand (children) or supply (parents). Having a brother might cause the daughter to demand more maternal and less paternal time. The results, however, clearly show that parents respond to sibling gender, which is the relevant margin, as any policy aiming at reducing the transmission of gender norms would most likely need to address parents and not children as young as 7 years.

[33] I observe the family structure on January 1^{st} each year and use the observation for the year the person turns 18 years or the last year in which the child lives with at least one biological parent.

[34] Not reported, considering heterogeneity by living in a traditional family for occupational choice shows that the effect is largest for women from divorced families. This is consistent with increased gender-specialization in these families. However, there is no significant heterogeneity by family structure for working in STEM occupations or choice of partner.

6 First-Born Men and their Second-Born Sisters

The main analysis investigates the effect of sibling gender on the origins of women's gender identity. This section briefly presents a corresponding analysis for men. However, I do not consider men's choice of partner or the school performance of their first-born children, because I find that sibling gender affects men's family formation both in terms of having a partner and having any children [Panel B in Appendix Table A12]; put differently, considering those outcomes might create selection issues and potentially bias the estimates. I construct the sample of men with identical selection criteria as for the main sample of women and conduct an identical analysis with the same variable definitions and controls.

Overall, the results for first-born men suggest that having a second-born sister relative to a second-born brother enhances men's gender identity [Appendix Table A11]. Men with a sister have a slightly higher (borderline significant) share of men in their occupation and are 0.51 percentage points (3.7 percent) more likely to work within STEM.[35] Importantly, however, having a sister also decreases the probability of working in managerial occupations by 0.44 percentage points (6.6 percent).[36] This decrease in the likelihood of working in (high-paid) managerial positions may help explain why men with a sister experience lower labor earnings than men with a brother [Appendix Figure A5]. At the same time, men with a sister cumulate less work experience at the end of their 30s relative to those with a brother, while there is no effect on lifetime unemployment by age 40. Thus, men with a sister appear somehow less successful in the labor market.

Similar to my findings, previous studies find negative effects of having sisters relative to brothers on men's earnings in Sweden and the U.S. (Peter et al., 2015; Rao and Chatterjee, 2017). Rao and Chatterjee (2017) show that in the U.S. brothers help each other more in job search than mixed sex siblings, which could help explain the negative effect on earnings and be a mechanism counteracting our ability to observe men's gender identity through occupational choice. Moreover, Peter et al. (2015) discuss competition between brothers as an important channel of the positive effect of having a brother on earnings. Brothers might compete with each other to a much

[35] The results are comparable when considering a binary indicator for having ever worked in STEM from age 31 through 40 (not reported).

[36] Not reported, I find a tight zero effect of sibling gender on women's probability of working in managerial occupations (the estimated effect is 0.04 percentage points ($se = 0.07$)).

greater extent than mixed sex siblings, both because men are more competitive than women and because having a same sex sibling might change the reference point of competition (Butcher and Case, 1994; Conley, 2000). Joensen and Nielsen (2017) show that especially brother pairs influence each other in terms of educational choice. Panel A in Appendix Table A12 shows that having a sister increases men's probability of ever enrolling in any field-specific STEM (traditionally heavily male-dominated) program, supporting a change in their gender identity. However, the effect does not persist into actual degree completion, which again may suggest that having a sister decreases competitive behavior, making them strive—and in the end—achieve less. Besides the effect on STEM enrollment, sibling gender does not impact men's educational attainment or achievement.

Like Peter et al. (2015), I also find that having a sister affects men's family formation negatively. Men with a sister cohabit and are married fewer years from age 18 through 41. Furthermore, having a sister decreases men's probability of having any children and their number of children. These findings could reflect less competitive behavior among men with a sister relative to those with a brother not only in the labor market but also in the marriage market. Thus, despite finding indications of similar effects of having an opposite sex sibling on men's development of gender norms as for women, competition might play a similarly or more important role for how men fare in the labor and marriage markets.

7 Conclusion

This study documents that the childhood family environment has a long-run impact on women's gender identity with persistent effects to the next generation of girls. The results show that having a second-born brother relative to a sister increases first-born women's gender conformity, both in terms of their choice of occupation and partner. I further show that having a brother negatively affects labor earnings. This is most likely driven by the effect on occupational choice, as sibling gender does not affect educational attainment, labor market participation, or family formation. I provide compelling evidence that changes in child-parent interactions—and, in particular, increased gender-specialized parenting in families with mixed sex children—play an important role for the changes in gender identity. This suggests that the transmission of traditional gender norms is stronger in families with mixed sex children. Finally, I show that the increased gender conformity among women with a brother persists into the next generation of girls, as indicated by an increase in daughters' comparative advantage in language over math performance in school. Consequently, I find evidence of very persistent long-run consequences of women's childhood family environment.

To eliminate gender inequality caused by gender-conforming behavior, my findings imply that policy makers need to focus on the formation of gender identity among girls in the childhood family environment. I show that having a brother affects girls' study choices in a more gender-stereotypical direction already at the end of compulsory schooling. This stresses that girls' development of gender identity by adolescence has important consequences for their later-life educational and labor market outcomes. As my mechanism analysis suggests, the family—representing a central aspect of the social environment during childhood—influences the formation of women's gender identity. Therefore, if society wants to give boys and girls the same opportunities at the time they enter the labor market in adulthood, policy makers would need to focus on how to counteract gender-stereotypical human capital investments. Specifically, interventions would need to counteract the transmission of gender norms across generations and thereby the development of gender-stereotypical behaviors, attitudes, and preferences.

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A Appendix

A.1 The Selection Bias Problem

To show the selection bias problem more formally, I here follow Peter et al. (2015). Assume a latent outcome $Y_i^* = \alpha + \beta G_i^{old} + X_i'\gamma + \epsilon_i$, where G_i^{old} is the gender of the older sibling and X_i is a vector of observable exogenous characteristics. ϵ_i contains other relevant unobservable variables, such as parental gender preferences denoted by P_i, and $E[\epsilon_i] = 0$. The bias arises because of the latent nature of Y_i^*, as we only observe the outcome if child i is born. In other words, $Y_i = Y_i^*$ if the child is born ($S_i = 1$) and Y_i is missing if the child is not born ($S_i = 0$). The selection depends both on parental preferences and the older child's gender, $S_i = f(P_i, G_i^{old})$. We can only estimate the effect for the sample of children who are born which gives the expected value of Y_i:

$$E[Y_i|S_i = 1, G_i^{old}, X_i] = \alpha + \beta G_i^{old} + \gamma X_i + E[\epsilon_i|S_i = 1, G_i^{old}, X_i] \tag{2}$$

$$= \alpha + \beta G_i^{old} + \gamma X_i + E[\epsilon_i|f(P_i, G_i^{old}) = 1, G_i^{old}, X_i].$$

As long as selection depends on the first child's gender and parental preferences affect the way in which parents raise their children $E[\epsilon_i|f(P_i, G_i^{old}) = 1, G_i^{old} = 1, X_i] \neq E[\epsilon_i|f(P_i, G_i^{old}) = 1, G_i^{old} = 0, X_i]$. This implies that the estimate of the older sibling's gender is biased.

A selection problem could also arise in the absence of parental gender preferences. Assume that first-born children have n normally-distributed traits, such as how easy the child is to take care of and how well it behaves. Suppose parents only want a second child if their first child has a value of each trait above a certain threshold. The threshold for or the distribution of each trait could be gender-specific. In both cases, parents who progress to the next parity would, on average, have different types of first-born children depending on the child's gender. For instance, if boys and girls have the same distribution of how well they behave but parents require girls to behave better than boys to have a second child, second-born children would, on average, have a better behaving older sibling if they have a sister compared to a brother. In this example, the estimated effect of the older sibling's gender on the younger child's outcomes might thus be due to the older sibling's behavior rather than due to his or her gender.

A.2 Family Size

Parents in developed countries are more likely to have a third child if their first two children are of same compared to mixed gender (Angrist and Evans, 1998; Angrist et al., 2010; Black et al., 2005). Appendix Table A1 shows that this is also the case in the main sample of the analysis. First-born women with a second-born brother are 13.2 percent less likely to have at least two siblings relative to those with a sister. The rest of this appendix examines whether family size has an independent effect on gender identity and studies rigorously the robustness of the main results to family size.

Table A1
Effect of Sibling Gender on Parental Realized Fertility

	# of Siblings (1)	≥ 2 Siblings (2)	≥ 3 Siblings (3)
Second-Born Brother	-0.07*** (0.01)	-5.26*** (0.28)	-1.33*** (0.16)
Observations		103,769	
Average	1.6	37.2	7.7

Estimates for the outcomes ≥ 2 *Siblings* and ≥ 3 *Siblings* are multiplied by 100 to express effects in percentage points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. *# of Siblings* measures the total number of siblings the individual has, including full and half siblings. $\geq 2(3)$ *Siblings* takes the value one if the person has at least two (three) full siblings and zero otherwise.

A.2.1 Does Family Size affect Gender Identity?

Black et al. (2005) use twins as an instrument for family size to show that family size does not affect educational attainment, using Norwegian registry data; Angrist et al. (2010) find the same for Israel. However, they only consider length of schooling and not gender identity. In this supplementary analysis, I show, consistent with their findings, employing a similar strategy in the Danish context, that family size does not affect educational attainment or the measures of gender identity used in the main analysis.

Table A2
The Effect of Family Size on Gender Conformity using Twins as Instrument

	First Stage	Second Stage				
		Choice of Occ and Partner			Education	
	# of Siblings	Log(Male Share in own Occ)	Works in STEM	Log(Female Share in Partner's Occ)	Log(Male Share in Edu)	Length (months)
	(1)	(2)	(3)	(4)	(5)	(6)
Twins at 2^{nd} parity	0.71***					
	(0.02)					
# of Siblings		3.82	0.70	-1.85	-1.33	0.27
		(3.35)	(0.82)	(4.84)	(3.78)	(1.06)
F-statistic of IV	1020.11					
Prob>F	< 0.001					
Observations	104,780	104,780	104,780	95,977	104,552	104,573
Effect×-0.07		-0.27	-0.05	0.13	0.09	-0.02

All second stage estimates (except Length of Education) are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample including twin siblings born at second parity (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. *# of Siblings* measures the total number of siblings the individual has, including full and half siblings. Columns (2) and (3) also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. Column (4) also includes dummies for the partner's number of occupational observations and age at first and last observation. The occupational outcomes of the first-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41. The effects are multiplied by -0.07 (*Effect×-0.07*), as it is the magnitude of the effect of having a brother on the number of siblings.

I use a sample with similar sample restrictions as for the main sample (see Subsection 3.1) with the exception that I include first-born singleton children who have younger twin siblings born at the second parity.[37] The instrument for family size is having twins at the second parity. Column (1) in Appendix Table A2 shows that the instrument is strong and relevant; see Angrist et al. (2010) and Black et al. (2005) for a discussion of the validity of the instrument.

Columns (2) through (6) show the second stage results. Similar to the findings for Norway and Israel, family size does not affect the length of highest completed education by age 30. Neither does it significantly impact the womans' occupational choice, her choice of partner, or her type of education. The last row in the table scales the estimates by -0.07 (i.e. the effect of having a second-born brother on the total number of siblings). This statistic (*Effect*$\times$*0.07*) illustrates that if family size (despite not having any statistically significant effect on the outcomes) would mediate some of the effect of sibling gender, any potential bias would be tiny.

A.2.2 Robustness to Family Size

As shown in Appendix Tables A1 and A3, sibling gender composition affects family size but family size does not affect gender identity. To further test the robustness of the main results to family size (in addition to flexibly control for family size as done in Column (4) in Table 3 in the main text), this subsection employs two alternative strategies: 1) to divide the sample by family size and 2) to study the effect of having a co-twin brother. Although family size is endogenous to sibling gender composition, strategy (1) is useful to the degree that it informs about the sensitivity of the results. These robustness analyses, together with the evidence of no differential effect by sibling gender on educational attainment or labor market participation [Table 5 and Figure 2] and the absence of an effect of family size on gender conformity, provide convincing evidence that family size does not confound the effects of sibling gender composition.

The first strategy is to split the sample by family size. For this, I restrict the sample to individuals who only have biological siblings, i.e. none of their parents have children with another person than the parent; though the results are similar when including those with half-siblings. Given family size is endogenous, this robustness check comes with a selection problem. If those parents of same sex

[37] I include all multiple births; twins, however, represent the vast majority of all multiple births.

Table A3
Splitting Sample by Family Size

	Log(Male Share in Own Occ)		Share of Years in STEM Occupation		Log(Female Share in Partner's Occ)	
	(1)	(2)	(3)	(4)	(5)	(6)
Second-Born Brother	-1.09*	-1.11	-0.49***	-0.37**	-2.26**	-1.84
	(0.62)	(0.82)	(0.16)	(0.18)	(0.88)	(1.12)
Observations	58313	36010	58313	36010	53148	33331
Average	788.4	784.9	5.5	4.4	299.3	290.7
# of Siblings	1	≥ 2	1	≥ 2	1	≥ 2

All estimates are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample with only full siblings (first-born women born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. Columns (1) through (4) also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. Columns (5) and (6) also includes dummies for the partner's number of occupational observations and age at first and last observation. The occupational outcomes of the first-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41. *1 Sibling*-models restrict the sample to those who only have one full sibling and no half-siblings. ≥ 2 *Siblings*-models restrict the sample to those who have at least two full siblings and no half-siblings.

children (born at the first two parities) who have a third child are more gender-stereotypical and to a greater extent influence their children's outcomes in such direction than those who do not have a third child, we would expect the effect of having a second-born brother to be larger in magnitude among first-born children from two-child families than for the entire sample. Similarly we would expect the effect of sibling gender to be smaller among children from families with at least three children. This is exactly what the results show in Table A3.

Table A4
Effect of Having A Co-Twin Brother on Gender Conformity in Education

	Next Birth	Log(Male Share in Edu)	STEM Enroll-ment	STEM Comple-tion
	(1)	(2)	(3)	(4)
Co-Twin Brother	-1.27*	-4.23**	-1.64***	-1.50***
	(0.73)	(2.04)	(0.58)	(0.43)
Observations	9,380	9,357	9,380	9,380
Average	28.9	331.7	7.3	4.2

All estimates are multiplied by 100 to express effects in percentage/log points. Standard errors in parentheses, clustered at the mother level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Each Column presents estimates from separate regressions. The sample consists of twins born 1962–86. All models absorb fixed effects for birth county, year of birth, mother's level and field of education, father's level and field of education, parity, and age at last educational observation. The models further control for (cubed) mother's age at birth and (cubed) father's age at birth. *Next Birth* indicates if the parents get a subsequent child. *Log(Male Share in Edu)* measures the natural logarithm of the male share in the highest completed education (narrow field-by-level) by age 30. *STEM Enrollment* indicates whether the woman has ever enrolled in a field-specific STEM education at age 16–27. *STEM Completion* indicates whether the woman has ever completed a field-specific STEM education by age 30.

Finally, to circumvent potential confounding effects from family size, I examine the effect of having a co-twin brother as an alternative empirical strategy. This approach is similar to the one in Cronqvist et al. (2015) and Peter et al. (2015), except that I do not have information on zygocity. To increase power, I include birth cohorts 1962–1986 and consider the gender conformity in educational outcomes. The key empirical feature of the sample of twins is that twin gender composition only has a very limited impact on family size [Appendix Table A4, Column (1)].

Overall, the effects of having a co-twin brother on educational choice are similar to the main results. The magnitude of the effects are, however, much larger. This may be due to the much greater intensity of the exposure to a co-twin compared to a younger sibling.

A.3 Educational System and Field of Study

Throughout, I follow the International Standard Classification of Education (ISCED) for the definition of all educational measures. I include observations through age 27 for all enrollment measures and through age 30 for all completion measures to give people time to complete the education in which they enroll. I define the *male share in education* as the share of men who had their highest completed education at age 30 within the same narrow field and level of education for cohorts born 1–5 years before the individual. The academic high school grade point average (GPA) is standardized with mean zero and standard deviation of one at the year of graduation and high school track level for the total population; note, however, that it is only observed for those completing the academic high school.

In the final year of 9^{th} grade, at age 16, students decide whether to apply for secondary education or to enter the labor market.[38] Secondary education (ISCED level 3) consists of two types: academic high school and vocational training. The academic high school is generic (i.e. not field-specific) and prepares students for tertiary education. For the cohorts of study, the academic high school had two tracks: language and math. Vocational education is, in contrast, field-specific and prepares students for specific occupations; I group Information and Communication Technologies and Engineering (ISCED fields 61 and 71) as STEM.

Tertiary education (ISCED levels 5–8) consists of three types: vocational, professional, and academic. I refer to the latter two jointly as *college*. Similarly, I group vocational secondary and vocational tertiary educations as *vocational education*. A vocational secondary degree usually only gives direct access to vocational tertiary programs within the same specific field,[39] while an academic high school diploma

[38] They can also choose to enroll in an optional 10^{th} grade, which is a formal a continuation of primary school. In the analysis, I restrict the attention to enrollment in and completion of programs after primary school, i.e. after grade 9 and 10.

[39] Students with a vocational secondary degree will often be required to have taken one or two academic high school courses at a basic level, such as Math and English. Many vocational secondary programs do not have a natural continuation at the tertiary level, though.

gives access to all types of tertiary education. An application to tertiary education is an application to a specific program. Most college STEM programs require certain high school STEM courses as prerequisites, such as advanced Math and intermediate Physics and Chemistry. Therefore, an academic high school STEM diploma gives much easier access to college STEM majors than other secondary school degrees, although it is possible to take complementary courses after high school graduation. Acceptance to college mainly depends on high school GPA and most STEM programs admit all eligible applicants (or have very low GPA cutoffs).

To mirror the definition of field-specific STEM education to the one of STEM occupation, I define STEM in college as Physical Sciences, Mathematics, Statistics, Economics, Information and Communication Technologies, and Engineering (ISCED fields 53, 54, 311, 61, 71). However, the results are similar when including Biology. Another important reason for excluding biology is that women's underrepresentation in STEM is limited to math-intensive —and, generally, better paid—science fields (Kahn and Ginther, 2017). The analysis of STEM education considers field-specific STEM educations in any type and at any level of education after primary school. This is to not potentially confound the results on STEM choice with educational attainment. Thus, the main STEM outcomes of interest indicate whether the individual ever enrolls in and completes a field-specific STEM education preparing for the labor market, including secondary and tertiary vocational STEM programs and college STEM majors.

Moreover, I complement the main STEM measures with four additional outcomes; the results are reported in the appendix. I examine whether the first place of enrollment after primary school has a STEM focus, i.e. whether it is either secondary STEM vocational education or in the math track in the academic high school. In line with this, I consider the probability of ever completing the academic high school math track. Finally, I split field-specific STEM educations by type, thereby investigating effects on the probability of completing a vocational STEM program and a college STEM major, separately.[40]

[40] Considering whether the highest completed education is within STEM reveals similar results as for having any field-specific STEM degree (not reported). Moreover, considering the probability of enrolling in the different types of STEM education rather than completing them also give similar results.

A.4 Alternative Mechanisms

This appendix describes alternative mechanisms to the ones discussed in Subsection 5.1. These mechanisms cannot be the dominating ones, however, as they are not compatible with the empirical findings.

The effect of sibling interactions might also go in the opposite direction for two reasons. First, the spillover model in developmental psychology hypothesizes that siblings imitate and influence each other with their gender-specific traits. For instance, Brim (1958) and Koch (1955) show that mixed sex siblings exhibit more traits of the opposite gender and fewer of their own gender compared to same sex sibling pairs. Second, the reference group theory in sociology suggests that as soon as a same sex sibling is present in the family, the same sex sibling will be the child and parents' reference group (Butcher and Case, 1994). Therefore, having a same sex sibling might induce the child to behave more gender-stereotypically. Meanwhile, given the empirical findings, neither of these two theories can be the dominating mechanism for the effect of sibling gender composition on the development of women's gender identity.

Studies examining the relationship between sibling gender composition and educational attainment have argued that budget constraints may play an important role (Amin, 2009; Butcher and Case, 1994). If parents face no borrowing constraints, they should, according to standard economic theory, invest in each child until marginal costs equal marginal benefits. However, if parents face borrowing constraints, they might decide to allocate their financial resources depending on the gender composition of their children. If parents want income equality between their children and the returns to education are smaller for women than men, then having a brother instead of a sister would be beneficial. However, parental aversion to income inequality cannot be the dominating channel, as we would otherwise have observed that having a sibling of the opposite sex should make the educational choice less gender-stereotypical.

In contrast, parents might want to maximize the total income of their children, thereby investing more in the child with the greatest returns to education. If returns to education are larger for men than women, having a brother would have adverse effects on educational attainment. In support of this argument, Powell and Steelman (1989) find for students enrolled in one college in the U.S. that the number of brothers puts more pressure on parents' financial support than do the number of sisters. Nevertheless, this is not a likely mechanism in the Danish context because there is no tuition fee at any educational level. Moreover, students in vocational training typically receive apprenticeship wages and students in tertiary education receive governmental student grants and loans to cover living expenses. For all cohorts in the analysis, students in tertiary education have at least had access to a combination of grants and loans of 1,000 USD a month in 2017-prices. It is also less clear how borrowing constraints should affect field choice, given sibling gender composition has no effect on the probability of enrolling in any type of program after compulsory education. Moreover, a more recent study shows that, for later generations in the U.S., parents to at least one son compared to parents with no sons do not differentially invest in their daughters (Cools and Patacchini, 2017).

A.5 Appendix Tables and Figures

Table A5

Effect of Sibling Gender on STEM Education and Educational Performance

	STEM Focus in First Enrollment	HS STEM Track Completion	Vocational STEM Completion	College STEM Completion	Grade 9 Language GPA	Grade 9 Math GPA
	(1)	(2)	(3)	(4)	(5)	(6)
Second-Born	-1.24***	-1.12***	-0.23**	-0.34***	-0.80	-0.80
Brother	(0.25)	(0.23)	(0.10)	(0.10)	(0.59)	(0.60)
Observations	103,769	103,769	103,769	103,769	82,978	82,350
Average	22.7	18.9	2.4	2.5	44.7	22.6

All estimates are multiplied by 100 to express effects in percentage points/percent of a standard deviation. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart) for STEM outcomes; girls born between 1986 and 1999 with the same selection criteria as for the main sample for the grade 9 outcomes. Each Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. The STEM outcome models further control for age at last observation in the education registry. *STEM Focus in First Enrollment* indicates whether the woman's first place of enrollment after compulsory schooling is in the academic high school math track or in a field-specific vocational STEM education. *HS STEM Track Completion* indicates whether the woman has completed the academic high school math track. *Vocational STEM Completion* indicates whether the woman has completed either secondary or tertiary vocational field-specific STEM education. *College STEM Completion* indicates whether the woman has completed a college degree or higher within STEM (excluding Biology). The Grade 9 GPA measures come from the written exam at the end of grade 9 in respectively Danish and Math and are standardized by year of graduation for the total population with mean zero and standard deviation of one.

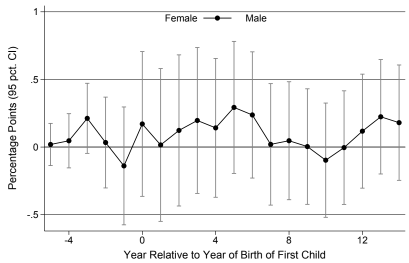
Table A6
Principal Component Analysis: Parental Time Investment

	Mother		Father	
	Age 7	Age 11	Age 7	Age 11
First Principal Component				
Play	0.51	0.58	0.49	0.53
Homework	0.32	0.37	0.47	0.43
Out-of-school activity	0.39	0.45	0.38	0.51
Read/sing	0.49	0.40	0.47	0.34
Excursion	0.49	0.42	0.41	0.40
Eigenvalue				
First Component	1.54	1.63	1.81	1.84
Second Component	0.97	1.09	0.92	0.95

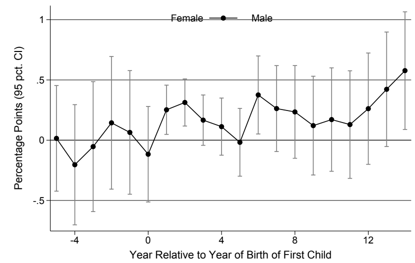
DALSC sample. Higher values reflect that parents do the specific activity more often.

Figure A1

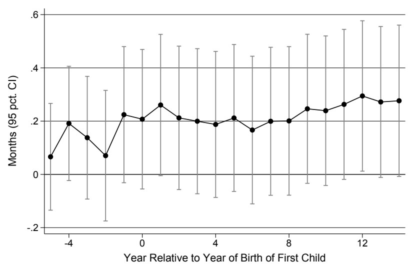
Parental Socio-Economic Status by Sibling Gender Composition



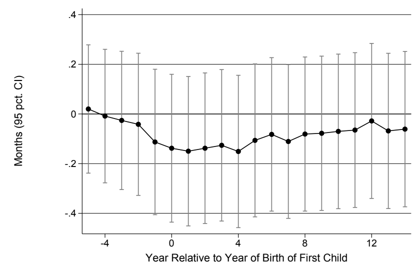
(a) Parents are Married



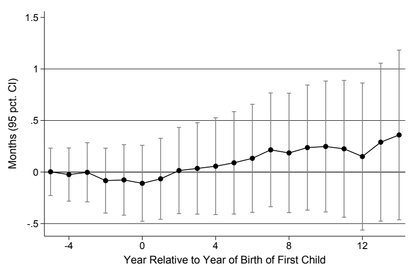
(b) Parents Cohabit/are Married



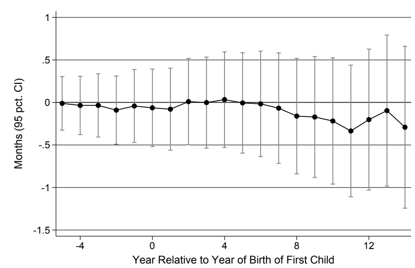
(c) Mother's Edu (months)



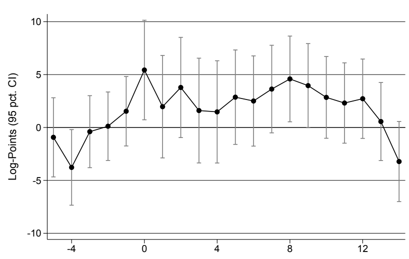
(d) Father's Edu (months)



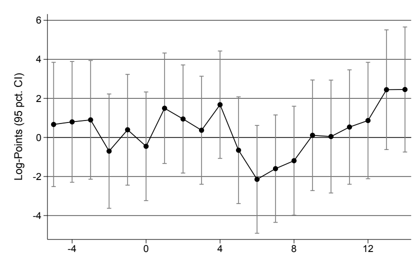
(e) Mother's Work Exp



(f) Father Work Exp



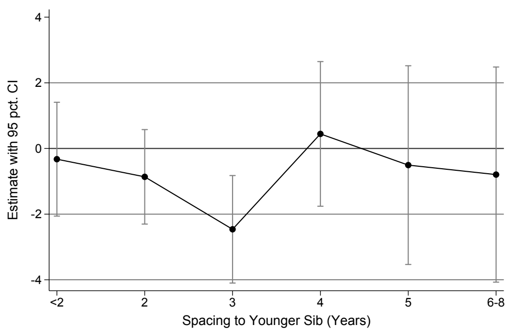
(g) Mother's log(Earnings)



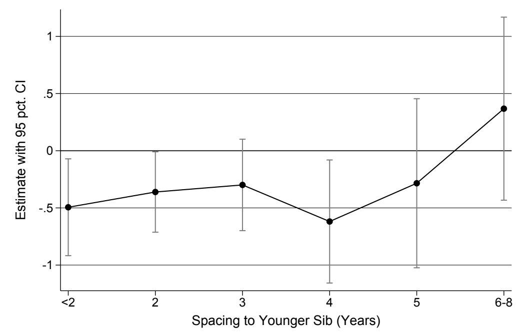
(h) Father's Log(Earnings)

Sample of first-born girls born between 1985 and 2002 with a second-born biological sibling born within four years apart. The whiskers represent the 95 percent confidence interval. All graphs illustrate the estimates from an event study of the effect of having a second-born brother. All models absorb time-specific fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education.

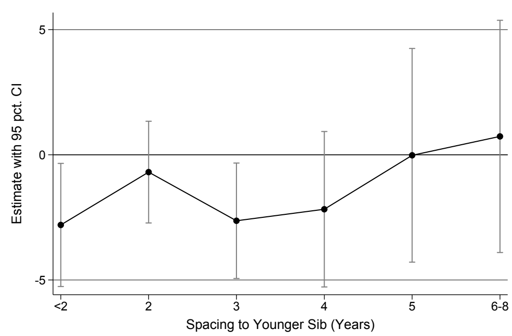
Figure A2
Effect of Sibling Gender on Choice of Occupation and Partner:
Heterogeneity by Spacing



(a) Log(Male Share in Own Occupation)



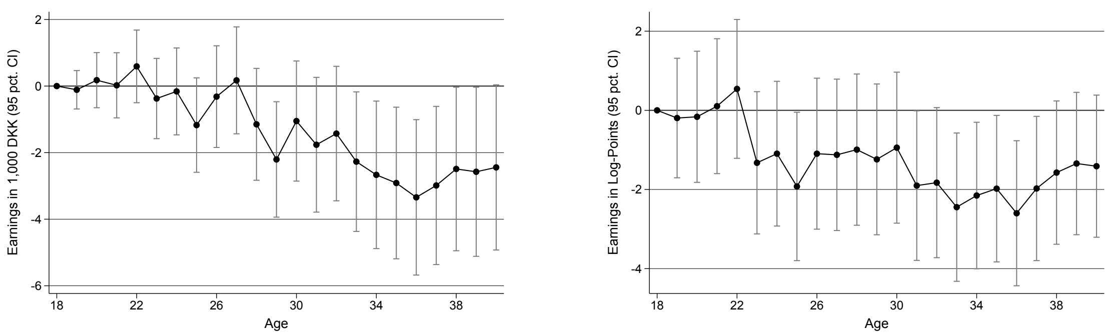
(b) Share of Years in STEM Occupation



(c) Log(Female Share in Partner's Occupation)

Main sample (first-born women born 1962–1975) including individuals with a second-born biological sibling born up to eight years apart. All graphs illustrate the estimated effect of having a second-born brother by birth spacing. The whiskers represent the 95 percent confidence interval. Each graph shows the estimates from a separate regression. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. The models with own occupation also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. For partner's occupation, the controls also include dummies for the partner's number of occupational observations and age at first and last observation. The occupational outcomes of the first-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41.

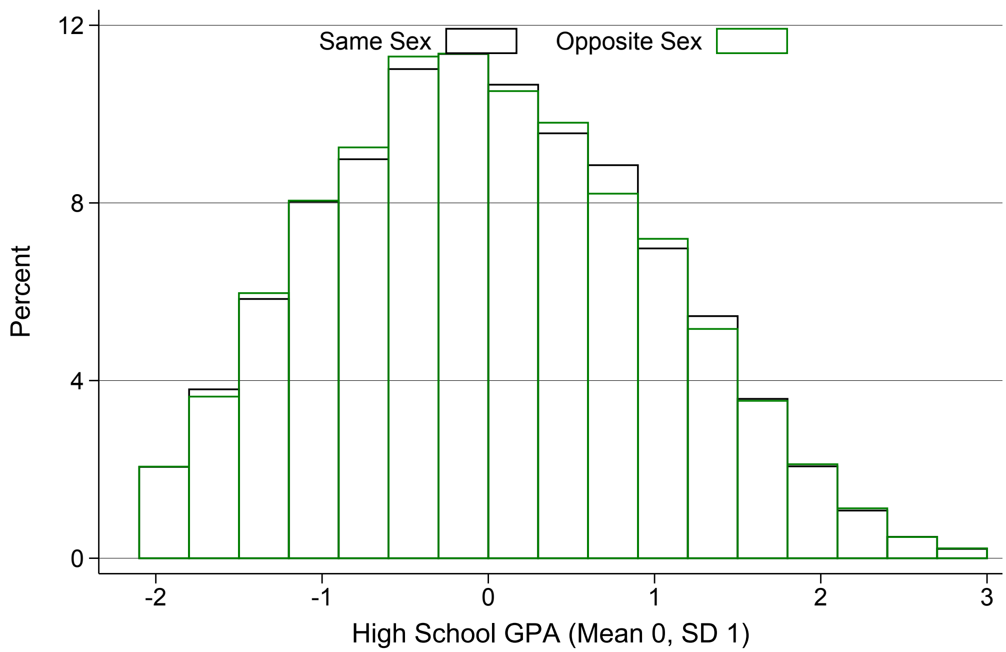
Figure A3
Effect of Sibling Gender on Earnings Age 18–40



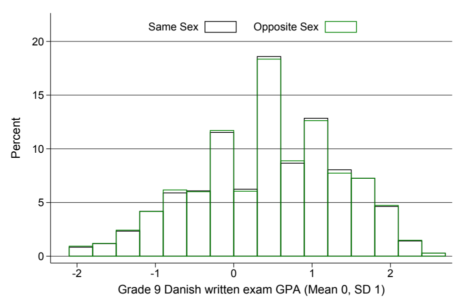
(a) Labor Earnings **(b)** Log(Earnings)

Main sample (first-born women born 1962–1975 with a younger biological sibling born within four years apart). The whiskers represent the 95 percent confidence interval. All graphs illustrate the estimates from an event study of the effect of having a second-born brother, where age 18 forms the base. Both models absorb time-specific fixed effects and individual fixed effects. *Labor Earnings* is measured in 1,000 DKK 2015-prices. *Log(Earnings)* is the natural logarithm of *Labor Earnings*.

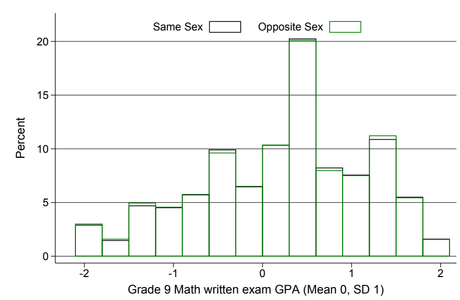
Figure A4
Distribution of Ability by Sibling Gender



(a) Academic High School GPA



(b) Grade 9 Language written exam



(c) Grade 9 Math written exam

Main sample (first-born women born 1962–1975 with a second-born biological sibling born within four years apart) for academic high school GPA; girls born between 1986 and 1999 with the same selection criteria as for the main sample for the grade 9 outcomes. The Grade 9 GPA measures come from the written exam at the end of grade 9 in respectively Danish and Math. *Academic High School GPA* is observed for students completing the academic high school language and math tracks. The standardized GPA measures are standardized by year of graduation (for the high school GPA track-by-year of graduation) for the total population with mean zero and standard deviation of one. All graphs plot the distribution of the three measures of school performance by gender of the second-born sibling [sister (black) and brother (green)]. The tails are truncated to have at least five observations within each cell due to data protection rules.

Table A7
Effect of Sibling Gender on Components of Parental Time Investment at Age 7 and 11

	Play	Home-work	Out-of-School Activ-ity	Read/Sing	Excur-sion
	(1)	(2)	(3)	(4)	(5)
Panel A: Maternal Investment at age 7 ($N = 594$)					
Second-Born	0.10	0.09	0.01	0.04	0.14*
Brother	(0.08)	(0.08)	(0.09)	(0.08)	(0.08)
Panel B: Maternal Investment at age 11 ($N = 594$)					
Second-Born	0.15*	0.14	0.01	0.11	0.09
Brother	(0.09)	(0.09)	(0.09)	(0.09)	(0.09)
Panel C: Paternal Investment at age 7 ($N = 421$)					
Second-Born	-0.10	-0.19*	-0.01	-0.28***	-0.00
Brother	(0.10)	(0.10)	(0.10)	(0.10)	(0.10)
Panel D: Paternal Investment at age 11 ($N = 415$)					
Second-Born	-0.16	-0.22**	-0.11	-0.14	-0.08
Brother	(0.10)	(0.10)	(0.10)	(0.10)	(0.10)

Standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. DALSC sample. Each Panel-Column represents the results from separate regressions. All models control for (quadratic) mother and father's age and fixed effects for spacing to the younger sibling in years, parental marital status in 1996, parents having been together for at least 5 years in 1996, region of birth, maternal level of education, paternal level of education, and family income level in 1995. Each of the individual components is standardized with mean zero and standard deviation of one.

Table A8

Effect of Sibling Gender on Quality of Child-Parent and Child-Sibling Relations

	Mother's	Father's	Child's relationship to		
	Relationship to Child		Mother	Father	Siblings
Child Age	11/15	7	15	15	15
	(1)	(2)	(3)	(4)	(5)
Second-Born	-0.12	-0.22**	0.07	-0.16*	-0.38***
Brother	(0.10)	(0.10)	(0.09)	(0.09)	(0.09)
Observations	441	434	498	489	485

Standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. DALSC sample. Each Column represents the results from separate regressions. All models control for (quadratic) mother and father's age and fixed effects for spacing to the younger sibling in years, parental marital status in 1996, parents having been together for at least 5 years in 1996, region of birth, maternal level of education, paternal level of education, and family income level in 1995. All child-parent relationship indexes represent the first component from principal component analyses, shown in Appendix Table A9, are standardized such that a higher value reflects a better relationship, the mean is zero, and the standard deviation is one. *Child's relationship to siblings* is an index of how easy the child thinks it is to talk to his/her siblings about matters that really bother her (standardized with mean zero and standard deviation of one).

Table A9
Principal Component Analysis: Child-Parent Relations

	Mother's	Father's	Child's rel'ship to	
	Rel'ship to Child		Mother	Father
First Principal Component				
Age 11: How close is the relationship between you and your daughter (1–4)?	0.71			
Age 15: How close is the relationship between you and your daughter (1–3)?	0.71			
Age 7: How close is the relationship between you and your daughter (1–4)?		0.71		
Age 7: Are you satisfied with the relationship between you and your daughter (1(yes)–2(no))?		0.71		
Age 15: Your mother/father plays a very big role in your life (1–5)			0.32	0.36
Age 15: Your relationship with your mother/father is important to you (1–5)			0.35	0.37
Age 15: Your mother/father loves you (1–5)			0.35	0.28
Age 15: You trust your mother/father (1–5)			0.38	0.40
Age 15: You can expect your mother/father to listen to you (1–5)			0.35	0.37
Age 15: You can go to your mother/father for advice (1–5)			0.40	0.36
Age 15: You can count on help from your mother/father if you have a problem (1–5)			0.36	0.37
Age 15: How easy is it to talk with your mother/father about matters that really bother you (1–5)			0.29	0.29
Eigenvalue				
First Component	1.34	1.25	4.07	4.53
Second Component	0.66	0.75	0.95	0.79

DALSC sample. All questions are answered on a likert scale with lower values being better. Therefore, the standardized measures used in Table A8 are all reversed, such that a higher value reflects a better relationship.

Table A10
Association Between First-Born Sibling's Gender and Second-Born Women's Gender Identity

	Log(Male Share in own Occ)	Works in STEM	Log(Female Share in Partner's Occ)	Log(Male Share in Edu)	Length (months)	High School GPA
	(1)	(2)	(3)	(4)	(5)	(6)
First-Born	-0.87*	-0.10	-1.20*	-1.15**	-0.42***	-0.05***
Brother	(0.46)	(0.11)	(0.67)	(0.52)	(0.15)	(0.01)
Observations	105,445	105,445	95,598	105,171	105,189	41,414
Average	787.6	4.623	292.2	333.6	156.9	0.002

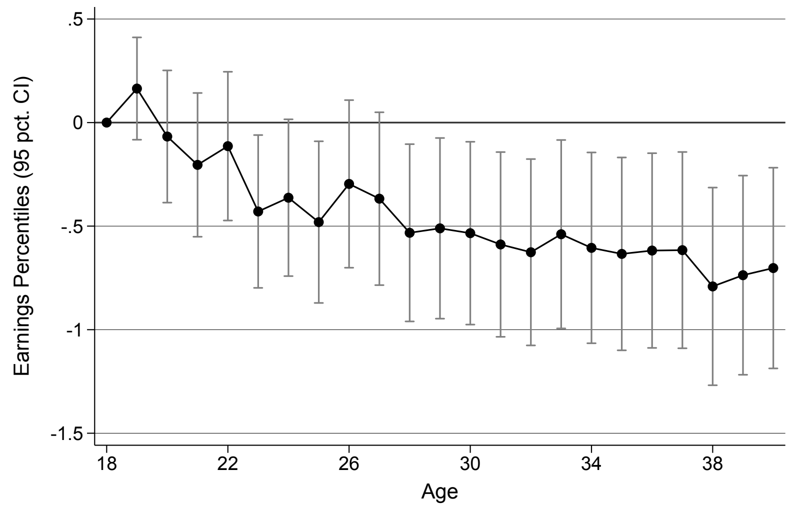
Sample of second-born women born 1962–1975 with a first-born biological sibling born within four years apart. Each Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to older sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. For the own occupation outcomes, basic controls also include dummies for the number of years observed in the income registry from age 31–40 and the number of years observed with a valid occupation code from age 31–40. For partner's occupation, basic controls also include dummies for the partner's number of occupational observations and age at first and last observation. The educational outcome models, except for high school GPA, further control for age at last observation in the education registry. The occupational outcomes of the second-born women are measured as mean from age 31–40. The occupational outcome of the partner is measured mainly at ages 31–45 for the partner with whom the woman lived most years from age 31–41. *Log(Male Share in Edu)* measures the natural logarithm of the share of men in the highest completed education (narrow field-by-level) by age 30. *Length* measures the length of the highest completed education in months by age 30. *HS GPA* measures final GPA from the academic high school and is standardized by track and year of graduation for the total population with mean zero and standard deviation of one.

Table A11
Men: Effect of Sibling Gender on Choice of Occupation

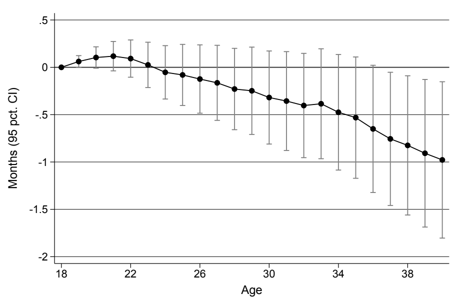
	(1)	(2)	(3)	(4)
Panel A: Log(Male Share in Own Occupation)				
Second-Born	0.51**	0.48*	0.44*	0.52**
Sister	(0.26)	(0.25)	(0.25)	(0.25)
Observations	108,366	108,365	108,365	108,365
Panel B: Share of Years Working in STEM Occupation				
Second-Born	0.44**	0.48***	0.51***	0.47***
Sister	(0.18)	(0.18)	(0.18)	(0.18)
Observations	108,366	108,365	108,365	108,365
Panel C: Share of Years Working as Manager				
Second-Born	-0.46***	-0.45***	-0.44***	-0.44***
Sister	(0.12)	(0.12)	(0.11)	(0.12)
Observations	108,366	108,365	108,365	108,365
No controls	✓			
Basic controls		✓	✓	✓
Parental education			✓	✓
Family size				✓

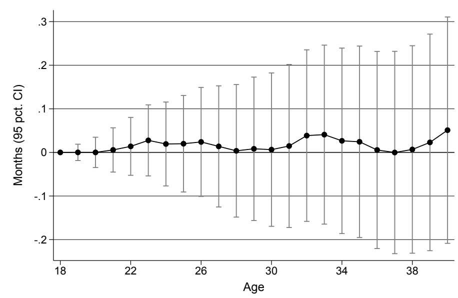
All estimates are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born *men* born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. *Basic controls* include fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, number of years observed in the income registry from age 31–40, and the number of years observed with a valid occupation code from age 31–40. *Parental education* controls include fixed effects for maternal level-by-field of education and paternal level-by-field of education. *Family size* controls include dummies for the number of biological siblings and dummies for the number of children the mother and father potentially have, respectively, from later relationships, and the gender of potential third- and fourth-born siblings. The outcomes are measured as mean from age 31–40.

Figure A5
Men: Effect of Sibling Gender on Labor Market Outcomes Age 18–40



(a) Earnings Percentile





(b) Cumulated Work Experience **(c)** Cumulated Unemployment

Main sample (first-born *men* born 1962–1975 with a younger biological sibling born within four years apart). The whiskers represent the 95 percent confidence interval. All graphs illustrate the estimates from an event study of the effect of having a second-born brother, where age 18 forms the base. All models absorb time-specific fixed effects and individual fixed effects. *Earnings Percentile* measures the labor earnings percentile by age and cohort. *Work Experience* measures the cumulated lifetime work experience in months. *Unemployment* measures the cumulated lifetime unemployment in months.

Table A12

Men: Effect of Sibling Gender on Education and Family Formation

	(1)	(2)	(3)	(4)	(5)
Panel A: Education by age 30					
	Log(Male Share)	Length (months)	High School GPA	STEM Enrollment	STEM Completion
Second-Born Sister	0.47	-0.06	0.01	1.16***	0.45
	(0.29)	(0.15)	(0.01)	(0.30)	(0.28)
Observations	107,898	107,921	31,973	108,365	108,365
Panel B: Family Formation by age 41					
	Cohabit 18–41	Married 18–41	Has Any Children	# of Children	Age at First Birth
Second-Born Sister	-0.39***	-0.88***	-1.56***	-0.04***	0.09***
	(0.12)	(0.15)	(0.24)	(0.01)	(0.03)
Observations	108,365	108,365	108,365	108,365	86,124

Estimates in Columns (1), (4), and (5) in Panel A and Columns (1), (2), and (3) in Panel B are multiplied by 100 to express effects in percentage/log-points. Standard errors in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Main sample (first-born *men* born 1962–1975 with a second-born biological sibling born within four years apart). Each Panel-Column presents estimates from separate regressions. All models absorb fixed effects for birth municipality, year-by-month of birth, spacing in months to younger sibling, maternal age at birth, paternal age at birth, maternal level-by-field of education, and paternal level-by-field of education. The educational outcome models (except for high school GPA), further control for age at last observation in the education registry. *Log(Male Share)* measures the natural logarithm of the share of men in the highest completed education (narrow field-by-level) by age 30. *Length* measures the length of the highest completed education in months by age 30. *High School GPA* measures final GPA from the academic high school and is standardized by track and year of graduation for the total population with mean zero and standard deviation of one. *STEM Enrollment* indicates whether the man has ever enrolled in a field-specific STEM education at age 16–27. *STEM Completion* indicates whether the man has ever completed a field-specific STEM education by age 30. *Cohabit* measures the share of years age 18–41 during which the man has cohabited with a partner without being married. *Married* measures the share of years age 18–41 during which the man has been married. *Has Any Children* indicates whether the man has at least one child by age 41. *# of Children* measures the number of children the man has by age 41. *Age at First Childbirth* measures the age at the man's first childbirth in years, conditional on having any children.

Chapter 2 – Exposure to More Female Peers Widens the Gender Gap in STEM Participation

Exposure to More Female Peers Widens the Gender Gap in STEM Participation*

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This paper investigates how high school gender composition affects students' participation in STEM college studies. Using Danish administrative data, we exploit idiosyncratic within-school variation in gender composition. We find that having a larger proportion of female peers reduces women's probability of enrolling in and graduating from STEM programs. Men's STEM participation increases with more female peers present. In the long run, women exposed to more female peers earn less because they (1) are less likely to work in STEM occupations, and (2) have more children. Our findings show that the school peer environment has lasting effects on occupational sorting and the gender wage gap.

(JEL I21, J16, J31)
Keywords: gender, peer effects, STEM studies

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I. Introduction

Although women today attain more education than men do, large gender differences in the choice of study field persist in most OECD countries. Only 28 percent of students in Science, Technology, Engineering, and Mathematics (STEM) studies are female (OECD, 2016). As gender differences in ability do not seem to explain these differences (Kahn and Ginther 2017), we currently know little about why women remain underrepresented in STEM fields. It is important to gain an improved understanding of the origins of gender differences in study choices due to the potential consequences for both the individual and society. First, women with high math and science ability who do not participate in STEM forfeit higher lifetime earnings. Second, society as a whole may be less innovative and thereby have worse long-run economic growth when fewer women are part of the STEM workforce.

In this paper, we investigate how the gender composition in high school affects men's and women's decisions to choose STEM fields in higher education. High school peers represent a central aspect of teenagers' social environment as they interact on a daily basis for several years. For students facing one of the most crucial life choices, peers may therefore represent an important social force shaping specialization decisions. To investigate whether gender composition in high school affects the gender gap in STEM participation, we use Danish register data on all students entering the math track in high school between 1980 and 1994. The key advantage of this data set, in addition to its rich information on individuals' education and labor market outcomes, is that we can follow the entire student population over a period of 20 years after they entered high school. This allows us to identify the direct, delayed, and long-run consequences of high school peers.

Our strategy to identify the causal impact of gender composition on STEM choice builds on two empirical approaches that differ in their key identifying assumptions and interpretation of results. The first empirical approach exploits idiosyncratic cohort variation in the proportion of female students *within schools* across cohorts after taking out school fixed effects, cohort fixed effects, and school-specific time trends.[1] The second empirical approach extends this model by including cohort-by-school fixed effects, thereby examining whether the proportion of female peers *differentially* affects men and women *within the same school-cohort*. This identification strategy thus answers whether the gender gap in STEM choice changes as the gender composition changes.

[1] This identification strategy is similar to Anelli and Peri (forthcoming), Hill (2017), Hoxby (2000), and Lavy and Schlosser (2011), who exploit idiosyncratic variation in the proportion of female students within schools.

The key identifying assumption for our first strategy is that year-to-year variations in the proportion of female students are exogenous to factors affecting STEM choice, conditional on school fixed effects, cohort fixed effects, and school-specific time trends. To assess the credibility of this identifying assumption, we conduct an extensive set of balancing checks, testing whether changes in gender composition are associated with student characteristics. Using a large set of student background characteristics from the register data, we show that gender composition does not systematically relate to the characteristics of students selecting into the specific school cohort, conditional on school and cohort fixed effects.[2] While this balancing test provides strong support for our key identifying assumption, it remains theoretically possible that students sort into schools based on factors correlated with STEM choice that are both time variant and unobservable in the register data. Although it is difficult to think of mechanisms that would create these unobservable time-variant school selection patterns, our second empirical approach addresses this concern. The inclusion of fixed effects for each cohort-by-school cell alleviates potentially remaining concerns, as we can control for the exact level at which selection based on *time variant* and *unobservable* characteristics would take place.[3] Both of our empirical approaches yield qualitatively similar results.

Our results show that women exposed to a higher proportion of female peers become less likely to enroll in STEM fields and more likely to enter health-related studies in college. Men also behave more gender-stereotypically when more female peers are present: they become more likely to enroll in STEM studies and less likely to enter health-related studies. These peer effects in field of study choice are statistically and economically significant. A 10-percentage point increase in the proportion of female high school peers lowers women's probability of enrolling in STEM studies by 1.4 percentage points—which is equivalent to a 7 percent decrease from the baseline. For men, a similar change in the gender composition raises STEM enrollment by 0.9 percentage points (2.3 percent). These peer effects exacerbate gender differences not only in STEM enrollment but also translate into an increased gender gap in STEM degree completion. In our most conservative

[2] We also test and reject the possibility that the proportion of female students enrolling in a given school is autocorrelated over time. Put differently, we find no evidence that the proportion of female students in year *t-1* predicts the proportion female peers in year *t*.

[3] Importantly, results from regressions including cohort-by-school fixed effects no longer identify whether an increase in the proportion of female peers affects the levels of STEM enrollment for men and women, but instead identify gender differences in the response to changes of the peer environment. To our knowledge, this is the first paper to apply this strategy to identify a gap in group-specific responses.

model, which includes cohort-by-schools fixed effects, we find that 10 percentage points more high school female peers increase the gender gap in STEM degree completion by 2 percentage points, corresponding to a 17 percent increase.

We shed some light on possible mechanisms behind this finding by studying how peer gender affects student performance, measured as high school grade point average (GPA), and how the effects differ across subgroups with different levels of parental education. Since high-achieving students are generally more likely to enter STEM fields and it has been shown that peers affect performance in high school (e.g. Hoxby, 2000; Lavy and Schlosser 2011), one possible mechanism is that the gender composition affects men's and women's preparedness for STEM studies. We find evidence in support of this mechanism: having more female peers alters the gender gap in high school GPA in favor of men, which may give women—who consider their comparative advantage—reason to believe that they are less prepared for STEM college studies. When considering heterogeneous effects by parental background, we provide two pieces of evidence that suggest that information about college, in general, and STEM studies, in particular, can counter peer influences to some degree. First, students with college-educated parents are less affected by peers. Second, and more strikingly, the gender peer composition does not influence women with STEM-educated mothers, i.e. women who have a salient female role model at home.

Our long-run results on labor market trajectories show that the peer effects in study choice lead women and men to systematically different career paths. Not only are women exposed to more female high school peers less likely to choose STEM studies, they are also less likely to work in STEM occupations and they have lower earnings at age 36. A 10-percentage point increase in the proportion of female high school peers lowers women's probability of working in STEM occupations by 6 percent and increases the gender wage gap by 5 percent. These results imply that high school peers and their influence on college major choice have lasting and economically significant consequences for occupational segregation and earnings.

To provide a better understanding of *why* we observe this sizable impact of peers on earnings, we examine fertility as an underlying mechanism. The relationship between high school peers and fertility is important for the interpretation of our results for at least two reasons. First, entering a STEM career may reduce fertility if jobs in these fields are less family-friendly. It might be harder to combine having children with work obligations in these environments. In this case, changes in fertility should be interpreted as an unintended consequence of the high school composition that ‘pushed’ some individuals in or out of STEM careers. Second, if peers directly affect preferences for children, then the documented effects for the gender wage gap may simply be because having children reduces earnings for women (Kleven et al. 2018; Lundborg et al. 2017). In this case, lower female earnings should not be attributed to staying out of STEM fields per se, but are instead a consequence of the "career cost of children" (Adda et. al., 2017).[4]

Notably, we find that women exposed to more female peers have more children by age 36. We provide suggestive evidence that about half of the effect of female peers on earnings is due to increased fertility and the other half is due to STEM participation. For men, we also find that peers affect fertility, but we find no impact on earnings. This is consistent with Kleven et al. (2018), which documents a large reduction in women’s earnings but no change in men’s earnings after the arrival of children.

In the existing literature, only a handful of related studies investigate how the gender of peers in high school influences educational choices.[5] Related to our work, Lavy and Schlosser (2011) find that both male and female students take more science courses in high school when exposed to a high school cohort with more female peers. While Lavy and Schlosser (2011) provide intriguing evidence on the underlying mechanism, it is not possible to study long-run effects on study choice, occupational sorting, fertility, or earnings in their setting. In contrast to Lavy and Schlosser (2011), Anelli and Peri (forthcoming) show that the gender composition in high school has an effect on men’s, but not on women’s, study choice in college. The authors find that men attending high school classes with over 80 percent male peers are more likely to enroll in predominantly male college majors. Contrary to our findings, these effects do not persist into actual degree completion or labor market outcomes. At the college level, Hill (2017) presents suggestive evidence that women exposed to a university cohort with more female peers have a lower probability of majoring in STEM fields. Similarly, Zölitz and Feld (2017) show that women become less likely to major in male-dominated subfields when they are randomly assigned to university sections containing more female peers. On the contrary, Schneeweis and Zweimüller (2012) show that a larger share of female peers in lower secondary vocational school increases

[4] Related to this point, if peers shape individuals' preferences for children—even before those children are born—men and women may choose more family-friendly careers outside STEM that can facilitate their family planning.
[5] Starting from Hoxby (2000) a different related strand of literature investigates how gender composition affects student performance. For important studies on the impact of peer gender on performance, see Whitmore (2005), Lavy and Schlosser (2011), Giorgi, Pellizzari and Woolston (2012), Oosterbeek and Ewijk (2014) as well as Hill (2015).

girls' propensity to choose male-dominated school types. Based on the existing literature, which presents mixed evidence from a variety of different settings, it is not clear how gender composition affects specialization decisions.[6]

This paper contributes to the literature in three important ways. First, we are the first to document that gender composition in high school affects STEM participation in college. Second, we provide previously undocumented comprehensive evidence on the long-run occupational consequences of high school peers. Our ability to follow students in the Danish administrative data over the course of 20 years after high school entry distinguishes this study from existing work that mostly studies the short- or medium-run impact of peers. Third, and more broadly, this paper contributes to a better understanding of the origins of gender differences in educational choices and labor market outcomes. This paper shows that the gender composition of high school peers represents an important aspect of the social environment that shapes individuals' preferences for field of study, occupation, and fertility and thereby influences earnings.

II. Institutional Background and Data

In this study, we use Danish administrative data covering the entire population of first-year high school students enrolled in the math track from 1980 through 1994. The key advantage of our dataset is that it contains rich background information and allows us to follow individuals over the course of 20 years after high school entry. We link the administrative data on high school students to annual data on educational enrollment and degree completion, which also contains detailed information on the type, level, and field of education, as well as labor market outcomes up to 20 years after entering high school.

Throughout this paper, we focus on students within the high school math track—students for whom entering STEM fields in college represents a relevant career option. As admission to most STEM college programs requires specific high school STEM course prerequisites, students from other high school tracks rarely choose STEM studies.[7] Among students within the math track, 30

[6] The related literature on the impact of single sex education also provides mixed results. While Jackson (2012) finds that single-sex secondary schools cause girls to take fewer math and science courses, other studies find no impact (Sohn, 2016) or a positive effect (Lee et al. (2014) and studies cited therein).
[7] For comparison, only 4.6 percent of STEM college graduates attended the high school language track.

percent later decide to enroll in a STEM program. These students thus represent the most relevant margin for increasing women's STEM participation. In the rest of this section, we introduce the institutional setting, describe the estimation sample, and present summary statistics on the key variables of the analysis.

A. *Institutional Background*

Children in Denmark enter primary school the year they turn 7 years and are required to attend school through grade 9.[8] It is optional to attend grade 10, which is a formal continuation of primary school. In their final year of primary school, students apply for secondary school. When applying for secondary education, students can choose between three-year academic high schools and vocational programs, which typically take four years.[9] The general academic high school, which represents the most popular type of academic high school, has two tracks: math and language.[10] In the high school application process, students specify their first, second, and third choice, each representing a combination of a specific high school and track. Students are qualified for high school admission if they have completed at least nine years of education with satisfactory results and if teachers state that they are qualified.[11] All applicants qualified according to these criteria are guaranteed admission in a high school in their county of residence. If there is insufficient capacity at all three preferred schools, the allocation committee in their home county admits them to another school after considering commuting time.[12] Schools experiencing capacity problems are concentrated in metropolitan areas. After high school completion, many students take one or two gap years before entering college.[13]

In the college application process, students apply for a specific field of study and a specific institution and can indicate up to eight institution-specific study programs. A diploma from an

[8] For the cohorts we study, it was not mandatory to attend a kindergarten class (grade 0), but most children did so.
[9] Academic high schools fall broadly into three branches: general, commercial (HHX), and technical (HTX).
[10] During the period we study, about 18 percent of each birth cohort enroll in the math track and about 45 percent of students within the math track are female. Both the share of math-track students and the share of women within the math track were relatively constant over the period our estimation sample covers.
[11] If these conditions do not hold, students can still qualify for school admission if they pass an entry exam.
[12] According to conversations with school principals active during our observation period, admission committees did not consider the gender of applicants during the admission process.
[13] We define college as professional and academic tertiary education.

academic high school is required for admission. Admission depends on high school GPA; however, most STEM programs have no or very low GPA cutoffs and almost all eligible students who apply are admitted. While GPA does not restrict students' STEM study choice, certain high school courses, such as advanced Mathematics and intermediate Physics and Chemistry, are prerequisites for STEM college majors.

B. *Estimation Sample*

We exclude students with missing values for gender and age (0.8 percent of students). We exclude students who were not between 14 and 19 years old when entering the general high school (less than 0.01 percent). We further restrict the estimation sample to schools in which at least 95 percent of students in a given cohort are 14–19 years old and exclude schools with very small cohort sizes of less than 10 students in a given year (6.1 percent of students). We apply these restrictions to exclude schools that mainly offer evening education or single courses, which target older part-time students who are, in many cases, working at the same time. Finally, we restrict the sample to schools that exist and admit students for at least four consecutive years (excluding 0.6 percent of students). None of these data restrictions qualitatively changes the results.[14]

C. *Summary Statistics*

Table 1 provides an overview of the summary statistics. Our estimation sample consists of 182,211 students attending 127 different schools over a period of 15 years, resulting in a total of 1,877 school-cohort observations. Forty-five percent of the students are female and the average cohort size is 108 students.

[Table 1 here]

[14] Results are available upon request.

Panel A in Table 1 shows the student outcomes we consider in this paper. The primary outcomes of interest are indicators for whether the student enrolls in a STEM study field and whether their college degree is within STEM fields at the college level or higher.[15] To classify STEM study programs, we follow the International Standard Classification of Education (ISCED) classification system. STEM degrees are thus studies within the following ISCED fields: Natural Sciences, Mathematics, and Statistics (ISCED-05), Information and Communication Technologies (ISCED-06), and Engineering, Manufacturing, and Construction (ISCED-07). To examine which fields within STEM drive the effects we also split STEM into four subfields: 1) Biology, 2) Math and Physics, 3) ICT and Engineering, and 4) Manufacturing and Construction. Additionally, we consider the probability of completing the highest degree within Health Sciences (ISCED-091), Education (ISCED-01), Arts/Humanities (ISCED-02), Social Sciences (ISCED-031), and Business/Law (ISCED-04).

From the total sample, 79 percent of all students enroll in college after high school. Table 1 shows that only 21 percent of female and 38 percent of male high school students subsequently enroll in STEM studies. This gender gap persists in STEM completion rates: while only 14 percent of women graduate with a STEM degree, 25 percent of men do so. Labor market outcomes show that 20 years after high school entry, 11 percent of women and 26 percent of men work in a STEM occupation.[16]

Panel B in Table 1 provides an overview of the student demographic and parental background characteristics we use in the regression analysis as controls, and Panel C shows school-level variables. The key peer variable of interest is the proportion of female peers at the time of high school entry, which we construct at the cohort-school-track level excluding the individual himself or herself. As less than one percent of students change to another high school or track, this group of peers represents the social group in which students interact over a three-year period. Students are, on average, exposed to 45 percent female peers. A one standard deviation change in the proportion of female peers is equivalent to 7.0 percentage points.

[15] Throughout the paper, we use the field of the highest obtained degree to construct measures of STEM completion. The enrollment variables are indicators for the student's ever having been enrolled in the respective study at the college level or higher. If we instead consider the field of first or last enrollment, we find very similar results throughout.

[16] We use the Danish version of the International Standard Classification of Occupations (DISCO) and construct an indicator for working in STEM if the individual works in a high-skilled occupation within STEM for at least half the years observed, 11–15 and 16—20 years after high school entry, respectively. All results remain qualitatively the same when using indicators for whether the mode occupation is within STEM for the considered periods.

Figure 1 shows the raw correlations between the proportion of female peers and the probability of completing a STEM degree. For women, a higher proportion of female peers is correlated with a lower probability of obtaining a STEM degree. For men, on the contrary, we observe a positive correlation between the proportion of female peers and STEM degree completion. These raw associations suggest a fairly linear relationship. While these correlations are purely descriptive, they foreshadow the results of our regression analysis.

[Figure 1 here]

III. Empirical Strategy

The fundamental threat to identification of peer effects arises from student sorting at various institutional levels. Parents select into neighborhoods, students select into schools, and within schools, students may select into classrooms or be assigned to tracks. As students are typically not assigned to schools at random, the existing peer effects studies try to overcome this identification problem by exploiting natural variation in cohort composition within a given school across time (Bifulco, Fletcher, and Ross 2011; Carrell and Hoekstra, 2010; Carrell, Hoekstra, and Kuka 2016; Hanushek et al. 2003; Hoxby 2000; Hoxby and Weingarth 2005; Lefgren 2004; Vigdor and Nechyba 2006). While this identification strategy addresses the issue of endogenous, time-constant, student sorting into schools, it is vulnerable to school-specific (dynamic) time trends that may alter both the peer composition and the outcome of interest. More recent peer effects studies respond to this concern with the inclusion of school-specific time trends—linear, quadratic, and cubic (Hill, 2017; Lavy and Schlosser 2011; Lavy, Schlosser, and Paserman 2012; Schneeweis and Zweimueller 2012). For identification, these studies exploit the deviation in peer composition from its long-term time trend within a school. This approach has the advantage of controlling for unobserved factors correlated with time trends in school composition that may confound peer effects in schools.

Our first empirical approach is similar to the approach in the literature discussed in the previous paragraph, and shares the key identifying assumption that the variation in the peer composition is exogenous after taking out school fixed effects, cohort fixed effects, and school-specific time trends. In our second empirical approach, we estimate an even more restrictive model by including fixed effects for each cohort-by-school cell, thereby alleviating all potentially remaining concerns regarding selection. In the following, we describe our empirical model, our two identification strategies, and the underlying assumptions in more detail.

A. *Empirical model*

Our main empirical model is:

$$Y_{isc} = \beta_1\, Female_i \times PropFemalePeers_{isc} + \beta_2\, Male_i \times PropFemalePeers_{isc}$$
$$+ \beta_3\, Female_i + C_{isc}\gamma' + e_{isc}\,, \tag{1}$$

where Y_{isc} is the outcome of student *i* attending school *s* in cohort *c*. The main outcomes we consider are STEM participation, the individuals' earnings percentile by age and birth cohort, and fertility. In our data, each individual represents one observation. The treatment variable of interest is $PropFemalePeers_{isc}$, which represents the proportion of female peers individual *i* is exposed to in their school *s* and cohort *c*. As the primary objective of this paper is to test whether peer composition affects women and men are differently, we interact the proportion of female peers with the indicator variables $Female_i$ and $Male_i$ that refer to the students' own gender.

β_1 thus captures to which degree women's study choice and labor market outcomes are affected by the peer gender composition in their high school and β_2 captures the equivalent impact for men. β_3 captures the gender gap in outcomes conditional on controls. C_{isc} represents a vector of school and cohort fixed effects as well as individual and peer characteristics which we gradually add when estimating Equation (1). The inclusion of high school fixed effects accounts for time-invariant endogenous sorting into schools and cohort fixed effects control for confounding factors at the national level, affecting all students in a given cohort. In order to account for unobserved time-variant school characteristics correlated both with changes in the proportion of female peers and educational choices for students within the same schools, we add school-specific linear, quadratic, and cubic time trends to the vector C_{isc}.

In our more conservative models, the vector C_{isc} includes the following additional student and peer average characteristics, which do not significantly alter our estimates: six indicator variables for mother's and father's highest educational degree and 18 indicators for their field of education;[17] indicators for first- and second-generation immigrant; a "traditional family" indicator that equals one if the student lives with both parents at age 10; dummies for student age at the time of high school start; mother's age at birth and its squared term; an indicator for having a young mother (< 22 years at birth); an indicator for whether the child is firstborn; family size and its squared term; and a dummy for whether the individual is adopted.[18] Additionally we control for up to third-degree polynomials of cohort size, as peer composition may potentially be correlated with cohort size (Epple and Romano, 2011).[19] Finally, the vector includes the proportion of female students in the language track in the same high school cohort and controls for the high school curriculum experiment that took place in Denmark in the 1980s (for more details regarding the experiment, see Joensen and Nielsen, 2016). The main purpose of including this large vector of control variables is to test how sensitive our results are to the inclusion of these variables. In the spirit of Altonji, Elder, and Taber (2005), changes in the coefficients of interest that result from including controls may inform us about the degree to which omitted unobservable factors may affect our results. To allow students' outcomes to correlate within their group of peers, we cluster the standard errors, e_{isc}, at the school-cohort level.[20]

The key identifying assumption for our first approach to yield causal estimates of β_1 and β_2 in Equation (1) is that no omitted variable exists that fulfills *all* of the following four requirements:

[17] For each parent, we include 9 dummies indicating whether their highest education is within ISCED fields 1–9.

[18] As individual controls, we additionally always control for dummies indicating the number of years after high school entry the individual was last observed in the education registry for our education models and dummies for the number of years observed within the given period for labor market outcomes. The inclusion of these controls reduces measurement errors given that the annual data does not record individuals living abroad.

[19] Moreover, we include indicators for whether the school experiences a change of more than 30 students or more than 50 percent in the cohort size compared to the previous cohort and up to two period lags of these variables. Given that a high school degree takes three years, we include these two-year lags to account for the possibility that a student's outcome was affected despite the fact that the inflow did not happen in their cohort. We furthermore control for the number of students from the cohort that are not in the age range of 14–19 years and for the number of students in the cohort that have missing gender and age information. Reasons for missing information on gender or age include that the person does not live in Denmark but attends a Danish school (e.g. lives close to the Danish-German border) or that the person resides in the country on Diplomat visa. None of these included control variables qualitatively changes our results.

[20] Angrist (2014) shows that with chance variation in peer groups, measurement error can bias peer effects estimates. Feld and Zölitz (2017) study this issue in more depth and show that classical measurement error can lead to overestimation of peer effects. Because we observe the students' gender in administrative registries, gender is arguably measured without error and our estimates should thus be free from upward bias arising from measurement error.

(1) time-variant and school specific,

(2) not captured by school-specific time trends,

(3) correlated with both the peer composition and the outcome of interest, and

(4) not included in the extensive set of individual- and peer-level control variables observed in the administrative data sets.

While it is difficult to think of any plausible mechanism that would create a violation of this type, the existence of such factors remains possible. To assess the credibility that such factors do not exist, we conduct an extensive set of balancing checks in section IV in which we test whether the peer composition in a school-cohort is systematically related to a large vector of high-quality measures of student background characteristics observable in the register data. While these balancing tests strongly support our key identifying assumption, we also provide results from a second empirical approach. Our second approach addresses the possibility of identification problems arising from unobservable, time-variant, and school-specific omitted variables not captured by school-specific time trends that may be correlated with both the peer composition and the outcome of interest.

In our second empirical approach, we extend Equation (1) by including an additional set of fixed effects for each cohort-by-school cell. The inclusion of these cohort-by-school fixed effects alleviates potential remaining concerns, as we control for the exact level at which selection based on *time variant* and *unobservable* characteristics would take place. Importantly, estimates from this type of model no longer identify whether an increase in the proportion of female peers will affect the overall number of women and men who choose a STEM program, but instead identify a gender difference in the response to changes in the peer environment. These estimates thus answer whether the proportion of female high school peers affects gender gaps in STEM participation and earnings.

IV. Balancing tests

To assess the plausibility of our key identifying assumption that *time variant* and *unobservable* factors are not driving our results, we test whether we observe systematic selection based on a wide range of observable student characteristics. One violation of our key identifying assumption would be, for example, if women select into a specific school based on the expectation of a higher or lower proportion of female peers within that school-cohort. In our first balancing test, we test whether students' own gender is correlated with the proportion of female peers, conditional on cohort and school fixed effects. This test closely follows the randomization check proposed by Guryan, Kroft, and Notowidigdo (2009) and controls for the school-level leave-out mean of the proportion of female peers across cohorts within the school to account for the mechanical relationship between own gender and peer gender. Table 2 shows that the proportion of female high school peers is not systematically related to students' own gender. The point estimate is precisely estimated and not distinguishable from zero. The inclusion of individual and school level controls as well as up to cubic time trends in Columns (2)–(5) does not significantly alter the point estimate.

[Table 2 here]

While Table 2 rejects sorting based on gender, it may still be possible that students sort into schools with a high proportion of female peers based on characteristics other than gender. The availability of a large set of high-quality measures of student background characteristics in the Danish administrative registries—typically not observable in other studies—allows us to rigorously test for this possibility.

In our next balancing test, we determine in how many cases student characteristics are significantly correlated with the proportion of female peers. Table 3 summarizes the significance of the point estimates from a total of 190 separate bivariate regressions, which test whether the proportion of female peers is related to student characteristics conditional on cohort and school fixed effects. Each column presents the results for a different set of control variables and school-specific time trends. Appendix Table A1 shows the full balancing test with all 190 coefficients. As expected when running a large number of regressions testing multiple hypotheses, some coefficients are statistically significant. In the absence of systematic sorting, we would expect 10 percent of coefficients to be statistically significant at the 10 percent significance level, 5 percent at the 5 percent level, and 1 percent at the 1 percent level simply due to chance. The share of significant coefficients is below the respective expectation for all three significance levels. Table 3 shows that 1.1 percent of estimates are significant at the 1 percent level threshold, 2.1 percent are significant at the 5 percent level, and that 8.9 percent are statistically significant at the 10 percent level. These balancing tests suggest that the proportion of female peers is as good as random and provide strong support for our key identification assumption. Without systematic cohort and school-specific sorting on this large set of observables, it appears highly unlikely that *unobservable time variant* factors create unobserved sorting patterns.

[Table 3 here]

Next, we test whether the proportion of female peers enrolling at a given school is autocorrelated over time. We do this by running 127 separate regressions that separately test, for each individual school, whether the proportion of female students at time t is correlated with the proportion of female peers that enrolled in $t-1$. While such school-level autocorrelation would not impose a threat to our identification strategy as it would be captured by the included school-specific time trends, the existence of such school-specific time dynamics may point to the existence of other unobservable time-variant confounders. Table 4 provides a summary of this exercise and reports the proportion of schools for which we find significant autocorrelation in the proportion of female students. For all significance levels, the share of schools for which the lag of female peers significantly predicts the proportion of female peers is close to what we would expect in the absence of autocorrelation. Across all models, 0.98 percent of the school-level regressions are significant at the 1 percent level, 3.35 percent are significant at the 5 percent level, and 8.86 percent are significant at the 10 percent level. Thus, we find no evidence that the proportion of female students enrolling in a given school is autocorrelated over time.

[Table 4 here]

As a final randomization check, we inspect whether the variation in the proportion of female peers, which we empirically exploit in this paper, is consistent with variation that we would expect with natural random fluctuations. Figure 2 plots the proportion of female peers at the school level after residualizing on cohort and school fixed effects and school-specific linear time trends. Figure 2 shows that these deviations in the proportion of female peers closely follow the normal distribution, which we plot for comparison. The shape of the distribution further supports the idea that the proportion of female peers is as good as random, conditional on the included controls.

In sum, the extensive set of balancing checks in this section provides strong support for our key identifying assumption. The evidence shown in Table 2, Table 3, Table 4, and Figure 2 suggests that the proportion of female peers is as good as random, conditional on cohort and school fixed effects.

V. Results

A. *Participation in STEM College Education*

Table 5 shows estimates of how the peer composition affects STEM enrollment (Panel A) and STEM degree completion (Panel B).[21] Column (1) shows the most basic model, which includes only the proportion of female peers, student gender, and the interaction between these variables. In Columns (2)–(6), we gradually include additional fixed effects and individual level controls. The specification in Column (6) includes school fixed effects, cohort fixed effects, a large set of peer- and student-level control variables, as well as linear, quadratic, and cubic school-specific time trends. Columns (2)–(6) show that the magnitude of the estimates is not particularly sensitive to the exact set of included fixed effects, controls, or time trends. Column (7) shows estimates from our most restrictive specification, which includes cohort-by-school fixed effects and shows the impact of peers on the gender gap in STEM participation.

[Table 5 here]

Our results show that women exposed to a higher proportion of female peers become less likely to enroll in and graduate from STEM college programs. Men's choices also become more gender-stereotypical: they are more likely to enroll in and complete STEM studies when they have a larger share of female high school peers. In our preferred specification in Column (4), a 10-

[21] We tested whether the gender composition of peers affects the probability of dropping out of high school and the probability of enrolling in or completing college. Table A2 in the Appendix shows no effect.

percentage point increase in the proportion of female high school peers lowers women's probability of enrolling in STEM by 1.4 percentage points, corresponding to a decrease of 6.7 percent relative to the baseline. For men, we find that a similar change in the gender composition raises STEM enrollment by 0.9 percentage points—a 2.4 percent change from the baseline.[22]

Column (7) shows our most restrictive model, which includes cohort-by-schools fixed effects. We find that 10 percentage points more female peers in a high school cohort increase the gender gap in STEM enrollment by 2.3 percentage points—which is equivalent to a 14 percent increase of the gender gap. Because we include cohort-by-school fixed effects in the model, the coefficient in Column (7) identifies a change in the gender gap in STEM completion and not an absolute effect. Importantly, the effect size we identify is close to the difference between the coefficients of male and female students in the less restrictive models in Columns (1)–(6), which increases our confidence in the estimates obtained from the models without cohort-by-school fixed effects.[23] Consequently, our estimate identified from within cohort variation implies that exposure to more female peers within a given school cohort substantially increases gender differences in STEM choice and leads to more gender-stereotypical enrollment choices. These results are consistent with Zölitz and Feld (2017), which find that exposure to more female peers in university teaching sections decreases women's likelihood of choosing male-dominated majors and lowers their labor earnings during the first years after college graduation.

Do high school peers only affect enrollment decisions or do they have lasting effects on study completion as well? The distinction between study enrollment and completion is potentially important as Anelli and Peri (forthcoming) find that gender peer effects in high school affect men's initial study enrollment, but have no impact on study completion or labor market outcomes. We therefore next shed light on the persistence of effects by considering the impact on the probability of STEM graduation.

Panel B in Table 5 shows that the peer effects in the field of enrollment persist into actual degree completion rates. In our setting, peer effects in study enrollment are not offset by changes of college major or college dropout. Women exposed to more female peers in high school are

[22] In addition to the linear-in-shares models shown in Table 5, we have also estimated non-linear peer effects using six bins for the proportion of female peers. In this analysis, we find relatively linear effects over the range of support that we have in the data (Figure 3 and Figure 4).

[23] To see this, compare the effect size of -0.265 in Column (7) with the estimate of -0.226 (-0.135 -0.091) in Column (4).

significantly less likely to graduate with a college STEM major. A 10-percentage point increase in the proportion of female peers lowers women's STEM graduation probability by 1.0 percentage points —a 7 percent decrease from the baseline (Column 4). The same change raises men's probability of graduating in a STEM field by 0.9 percentage points, which is equivalent to a 3.6 percent increase from the baseline. Again, the point estimates of interest in Columns (2)–(6) are very similar across models and are insensitive to the exact set of included fixed effects and time trends. Column (7) in Panel B confirms that these results hold when including cohort-by-school fixed effects. Gender differences also remain present in graduation rates when we exploit whether the gender composition among peers differentially affects men and women within the same school-cohort. Column (7) shows that 10 percentage points more female peers increase the gender gap in STEM completion by 2 percentage points, corresponding to 17 percent.[24]

Given the low baseline rates for women's STEM enrolment and graduation, the size of the peer effects we document in Table 5 are economically significant. Taken together, we find that a higher share of female peers makes both men's and women's initial choice of study field and field of graduation more gender-stereotypical.

To understand which STEM subfields women are less attracted to when more female peers are present, we next estimate separate models in which we split STEM into four subgroups, shown in Panel A of Table 6: (1) Biology, (2) Math and Physics, (3) ICT and Engineering and (4) Manufacturing and Construction. . A comparison of the point estimates in Columns (1)–(4) reveals that the coefficients for women are relatively similar across STEM subfields. The point estimate for ICT & Engineering is marginally smaller and less precisely estimated, but not significantly different, from the other subfields. For men, the effect of having more female peers is stronger for entering Math and Physics and ICT and Engineering, which are the most male-dominated, gender-stereotypical STEM subfields.

We next ask which other study fields become more attractive for women and less attractive

[24] As a robustness check, we also test whether results differ between students who attend a high school that is one of several in the municipality and those who attend the only high school in the municipality. If our estimates were driven by unobserved, time-variant selection into high schools, we would expect effects to differ substantially based on how much choice students have at the local level. Table A3 in the Appendix reports split sample regressions by the number of high schools in the municipality. For women, point estimates are very similar in regions that have only one high school in the municipality. For men, we find that the effect of peers seems to be somewhat larger in municipalities with only one high school. While it is possible that peer effects may differ by the number of high schools in the municipality, we think that these results provide additional support for the validity of our peer effects estimates as we find the same effects for regions where students had a de facto only very limited school choice.

for men when having experienced a larger share of female high school peers. In Panel B of Table 6, we examine how the peer composition affects graduation from (1) Health Sciences, (2) Education Studies, (3) Arts and Humanities, (4), Social Sciences, and (5) Business and Law. Only Health Sciences—a field heavily dominated by women—becomes significantly more popular among women who had more female peers in high school (Column 3).[25] For men, we also find that more female peers make choices more gender-stereotypical: a larger share of female peers in high school decreases men's probability of graduation with a college major within Health and Education Sciences. Columns (2)–(5) show that the high school gender composition does not influence women's probability of completing a college degree in Education, Arts and Humanities, Social Sciences, or Business and Law. Consequently, our results suggest that women exposed to more female peers in high school substitute STEM studies with careers in the health sector. In Subsection C, we investigate the potential labor market consequences of these education choices in more detail.

[Table 6 here]

B. *Underlying Mechanisms and Heterogeneity*

Why does the gender composition in high school affect students' decision to enter STEM fields? We shed some light on the underlying mechanism in two ways. First, we investigate whether the high school gender composition affects final high school GPA, which students use to apply for college. Second, we split the sample based on parental educational level and field of education to learn more about whether some groups of students are more sensitive than others to the gender composition at their school. Considering heterogeneity in a parent's field of education might help us understand whether STEM role models at home moderate or perpetuate the influence of peers.

In order to assess the first proposed mechanism, we investigate whether gender composition directly affects students' study ability or preparedness to enter STEM studies. This appears plausible as the peer effects literature has shown that gender composition can impact students' performance (Hill, 2017; Hoxby, 2000; Lavy and Schlosser, 2011). If the gender composition

[25] Of all women who enter health sciences 50 percent study nursing and midwifery, 20 percent study medicine, and 13 percent therapy and rehabilitation.

affects student performance differentially depending on the student's own gender, this effect may in part rationalize the effects on the choice of college major. Table 7 provides estimation results supporting such a mechanism.

Column (1) in Table 7 shows that the gender composition does not affect women's GPAs; the point estimate is tiny and not statistically significant. On the contrary, male students achieve a higher GPA when they are exposed to a high school cohort with more female peers. Ten percentage points more female peers raises the GPA of male students by 1.26 percent of a standard deviation.[26] Our finding is consistent with Lavy & Schlosser (2011) and Hill (2017), which show that men in high school and college achieve better grades when there are more female peers in their cohort. Importantly, the vast majority of STEM college programs in Denmark do not have a binding high school GPA threshold for admission. This rules out the possibility that gender composition mechanically affects STEM enrollment through the impact on male students' GPA.

The fact that male students achieve a higher GPA when a higher proportion of female peers are present may in part explain why fewer women and more men enter STEM studies in cohorts with more women. Given their higher GPAs, men might feel better prepared for STEM studies, which generally attract students with better high school grades. In contrast, women do not perform differently in high school, but might infer from the gender gap in high school GPA that they are less prepared or "suited" to enter STEM studies than their male peers. These results are consistent with Zölitz and Feld (2017), who also find evidence of gender-specific performance responses that can rationalize students' specialization choices.

Table 7 further investigates whether the influence of female peers is similar for students with less- versus highly educated parents (Columns 3 and 4) and parents with STEM educations (Columns 5 and 6). Our motivation for splitting the sample by parental education is that students from nonacademic families may have less information about college majors and associated occupations and therefore be more sensitive to their peers' choices. If peers can provide information about study fields that is not available to students who have non-college educated parents, we would expect students from families with less-educated parents to be more sensitive to the peer composition. Similarly, students with a STEM-educated father or mother might have better information about STEM studies and careers. If parents serve as strong role models that shape the

[26] All results in Table 7 are robust to the inclusion of cohort-by-school fixed effects.

choices of their children, we would expect that students with a (same sex) parent in STEM are less sensitive to peer influences.

Column (3) shows estimates for the subsample of students who have parents without a college education, while Column (4) shows the same model for students where one or both parents have a college degree.[27] The results show that the influence of peers is twice as strong for women with parents without college education relative to women with college-educated parents. However, effects for men do not vary by parental education. One interpretation of these results is that women are more likely to follow the choices of their peers when they lack parents who have a greater capacity to help them find information about higher education or share their own college experiences. It is also possible that the study choice of parents who attended college provides an additional reference point that moderates the impact of peer effects in high school.

We next investigate whether having a father or mother with a STEM education mediates effects of the peer composition. Column (5) shows that the effect of peers on STEM completion is similar for women with a father in STEM and somewhat weaker and not statistically significant for men with fathers in STEM. These results suggest that men who have a STEM father are less susceptible to peer influences. Column (6) shows that these results are mirrored for women with STEM mothers. Strikingly, women with STEM mothers are not significantly affected by the peer gender composition. The point estimate is in fact positive, which suggests that STEM mothers counteract the effect of peers on their daughters' specialization choice. While the group of individuals who have a mother in a STEM field is small, these results show that women who have a STEM mother as a role model are unaffected by peers. While these results remain suggestive, they could imply that access to a non-stereotypical same sex role model in the family is more powerful in shaping women's STEM interest than the influence of high school peers.

[Table 7 here]

C. *Long-Run Effects on Labor Market Outcomes*

[27] To facilitate comparison, Column (2) in Table 7 reports point estimates for the full sample from our preferred specification in Column (5) of Table 5.

Given our finding that high school gender composition affects the probability of enrolling in a STEM college major and completing such education, we now want to know whether these gender peer effects also persist into occupational choice and whether they have consequences for labor earnings.

[Table 8 here]

Table 8 shows the impact of high school peers on labor market outcomes 15 and 20 years after high school entry (Columns 1 and 2). Fifteen years after high school entry, the median age of individuals is 29 years and most should have completed higher education and entered the labor market.[28] Table 8 shows that peers' influence on STEM field choice are persistent and closely mapped into women's occupational choices. A larger share of female high school peers reduces the probability that women work in a STEM-related occupation, while it has no impact on men's occupational choice. A 10-percentage-point increase in the proportion of female peers decreases the probability that women work in a STEM occupation 20 years after high school entry by 0.50 percentage points—a 4.6 percent decrease from the baseline (Column 2). In contrast, men's probability of working within STEM fields is not affected by the peer composition. This suggests that, although men's study choice is affected, those who enroll in and graduate from STEM fields due to peer effects are not more or less likely to end up working in STEM jobs.

Given these long-run effects on occupational choice, we next test whether individuals' labor earnings are affected. Does the proportion of female high school peers contribute to the gender wage gap? Columns (3) and (4) shed light on this question and show estimates for earnings 15 and 20 years after high school entry. We find that the high school gender composition has lasting effects on women's but not on men's earnings. A 10-percentage-point increase in the proportion of female peers decreases women's earnings 15 years after high school entry by 0.39 percentile points—nearly a one-percent decline from the baseline—and has no impact on men's earnings. Twenty years after high school entry, we still do not find any effect for men but an even larger effect of 0.67 for women. This increase in magnitude for women is in line with a similar increase in the

[28] Brenøe and Lundberg (2017) show that after age 30, the share of a cohort that has completed a college or university degree is almost constant, indicating that by age 31 most individuals have completed higher education.

effect on their STEM occupational participation. The estimate from our second strategy (Panel B), which includes cohort-by-school fixed effects, suggests that having a 10-percentage-point-larger share of female high school peers increases the gender earnings gap by 0.75 percentile points, corresponding to an increase of 5.4 percent (Column 4). Thus, our results show that high school gender composition has lasting impacts on the gender wage gap, likely mediated by the impact on STEM participation.

D. *Long-Run Effects on Fertility*

Given that we find a substantial effect of the high school gender composition on women's earnings, we next examine fertility as an underlying mechanism. While fertility is an interesting outcome in itself, we believe that there are a least two reasons the relationship between high school peers and fertility is important for the interpretation of our earnings results. First, women entering a STEM career because of peers may have fewer children once they realize that STEM jobs are less family-friendly. The more competitive environment, longer working hours, and lower job flexibility might make it harder to combine children with work obligations. In this case, changes in fertility should be interpreted as an unintended consequence of the high school composition that "pushed" some individuals in or out of STEM careers. Second, if the high school gender composition directly affects fertility preferences, then the documented effects for the gender wage gap may simply be due to the fact that having children reduces earnings for women (Kleven et al. 2018; Lundborg et al. 2017). In this case, lower earnings for women should not be attributed to their staying out of STEM fields per se, but may instead be a consequence of the "career cost of children" (Adda et. al., 2017).

[Table 9 here]

Table 9 documents the medium- and long-run consequence of high school gender composition on fertility. Columns (1)–(4) provide a detailed analysis on the timing of fertility effects on the extensive margin. While Column (1) shows little effect on fertility during the first 5 years after high school entry, we see a clear impact on the probability of having any children within

10 years after high school entry when individuals are around 26 years old and for the most part have entered the labor market (Column 2). We find that having a larger proportion of female high school peers increases women's and decreases men's probability of having any children by the age of 26. Column (4) shows that 20 years after high school entry, women are no longer less likely to have any children, but we still find a persistent effect for men.[29]

In Columns (5)–(8) of Table 9, we estimate effects at the intensive fertility margin by testing whether female peers affect the number of children individuals have. We find that women, who had more female peers and thus were less likely to enter STEM careers, have more children by the age of 36. The increased fertility effect becomes visible shortly after the time of college completion and doubles in size within the first couple of years in the labor market (Column 7). By the time individuals are 36 years old, women exposed to 10-percentage-points more female peers have on average 0.02 more children, an increase of about 1.0 percent from the baseline of 1.66 children (Column 8). For men, we find that those exposed to more female peers have significantly fewer children by age 36.

Our results show that the proportion of female peers has not only significant impacts on study choice, occupational choice, and earnings, but affects fertility as well. Strikingly, a larger share of female high school peers increases women's number of children. If high school peers already affect men's and women's fertility preferences before choosing their field of study, women's shift from STEM to health-related studies (and the reverse for men) might partly be explained by changes in the desired number of children.

These findings on fertility raise the question of whether the negative effect of having more female high school peers on women's earnings documented in Table 8 is driven by: (1) women's not participating in STEM, or by (2) their increased fertility. Although we cannot provide causal evidence that distinguishes between these mechanisms, we still attempt to provide some suggestive evidence. In Table 10, we test whether we can reduce or eliminate the estimated effect of the proportion of female peers on the earnings by including controls for fertility and STEM participation. While the inclusion of these endogenous variables, commonly referred to as "bad controls," complicates the interpretation, we still believe that this exercise hints at whether the earnings results can exclusively be attributed to gender peer effects on STEM participation or

[29] Restricting the sample to older cohorts reveals very similar effects on fertility through older ages. Therefore, the reported affects are close to effects on complete fertility, especially for women.

fertility. Column (1) shows the estimated main effects on earnings we observed in Column 4 of Table 8. In Column (2), we add controls for the individual's fertility by including gender-specific dummies for the number of children of the individual. In Column (3), we do not control for observed fertility but instead we control for STEM participation using gender-specific dummies for working in a STEM occupation and having a STEM college degree, and an interaction term between STEM occupation and education. Finally, Column (4) includes both the fertility and STEM controls.

[Table 10 here]

The results in Table 10 suggest that including either the fertility or STEM controls reduces the estimated peer effect on women's earnings and the gender gap in earnings by 40–50 percent. Once both sets of controls are included, the estimated gender peer effect is no longer significant and is only 21 percent of its original magnitude in Panel A. These findings indicate that the effects of the high school gender peer composition on fertility and STEM participation equally contribute to the long-run consequences for women's earnings. However, given the "bad control" problem we face here, we caution a causal interpretation of this result.

These results suggest that fertility cannot fully explain the lower earnings of women who were exposed to more female peers. These results are consistent with Kleven et al. (2018), who use Danish administrative data (similar to our data) and document that men and women experience very similar trends in labor earnings before the arrival of their first child. Yet, Kleven et al. (2018) show while women experience a large decline in earnings at the time of their first childbirth and still have earnings 20 percent below their initial level ten years later, men do not experience any change in their earnings after having children.

Taken together, our results suggest that there are two key factors to explain why women with more exposure to female peers have lower earnings. First, entering STEM jobs, which are arguably more competitive and have less flexible working hours, reduces women's fertility. Second, exposure to more female peers in high school increases women's fertility, which reduces their earnings independently of their STEM participation.

VI. Conclusion

This study demonstrates that the gender composition of high school peers affects students' decisions to undertake STEM studies in higher education. Our results show that a higher proportion of female high school peers makes study choices more gender-stereotypical. With more female peers present, women become less likely to enter STEM fields and more likely to enter Health Studies. Men also behave more gender-stereotypical and become more prone to enter STEM studies when exposed to more female peers.

For women, these gender peer effects in study choice have remarkably persistent long run effects on occupational choice, which remain visible 20 years after high school entry. Women who by chance were exposed to more female peers are less likely to work in STEM occupations and have lower earnings 20 years after high school entry. We further show that the high school gender composition affects individuals' fertility and provide suggestive evidence that gender peer effects on STEM participation and fertility equally contribute to the effect on women's earnings.

In conclusion, our results suggest that gender peer effects in high school shape preferences for study fields and thereby lead students to systematically different career trajectories. Our evidence on the underlying mechanisms remains suggestive, but indicates that a higher proportion of female peers affects the gender gap in high school GPA and may therefore foster the gender gap in STEM preparedness, which gives students reason to believe that they have a comparative advantage in a more gender-stereotypical college major. We also find that women with STEM-educated mothers are unaffected by the gender composition, which suggests that salient female role models may be able to counteract peer pressures in high school.

We believe this paper broadens our understanding of where gender differences in educational choice originate. Our findings emphasize that the social environment directly affects students' decisions to specialize within STEM fields and educational and occupational choices more generally. Moreover, it highlights the possibility that manipulating the gender composition in a given environment through affirmative action policies to achieve gender balance may have adverse and unintended consequences for fertility, gender segregation in college majors, and the labor market.

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TABLES

Table 1: Descriptive Statistics

	Women			Men		
	N	Mean	SD	N	Mean	SD
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Student outcome variables						
Any college enrollment	81820	0.794	0.404	100391	0.784	0.411
Any STEM enrollment	81820	0.210	0.407	100391	0.378	0.485
Any Science/Math enrollment	81820	0.118	0.323	100391	0.131	0.338
Any Technology/Engineering enrollment	81820	0.104	0.306	100391	0.279	0.448
Any Health enrollment	81820	0.268	0.443	100391	0.063	0.242
Any college completion	81820	0.710	0.454	100391	0.650	0.477
Highest completed degree within STEM	81820	0.135	0.342	100391	0.253	0.435
Highest completed degree within Science/Math	81820	0.062	0.241	100391	0.056	0.229
Highest completed degree within Technology/Engineering	81820	0.073	0.260	100391	0.197	0.398
Highest completed degree within Health	81820	0.217	0.412	100391	0.047	0.211
Highest completed degree within Health	81820	0.217	0.412	100391	0.047	0.211
Highest completed degree within Education	81820	0.075	0.264	100391	0.040	0.197
Highest completed degree within Arts & Humanities	81820	0.058	0.234	100391	0.045	0.208
Highest completed degree within Social Sciences	81820	0.049	0.215	100391	0.062	0.242
Highest completed degree within Business, Admin, Law	81820	0.093	0.290	100391	0.134	0.341
STEM occupation 15 years after high school entry	81270	0.103	0.304	99752	0.246	0.431
STEM occupation 20 years after high school entry	80114	0.109	0.312	98126	0.255	0.436
Annual labor earnings 15 years after high school entry	81265	241.944	133.320	99747	311.796	187.452
Annual labor earnings 20 years after high school entry	80108	327.915	167.152	98121	454.849	272.011
Earnings percentile by cohort 15 years after HS entry	81265	52.2	22.9	99747	62.6	26.3
Earnings percentile by cohort 20 years after HS entry	80108	56.6	23.7	98121	70.6	26
Any children 5 years after HS entry	81820	0.019	0.136	100391	0.007	0.081
Any children 10 years after HS entry	81820	0.208	0.406	100391	0.108	0.310
Any children 15 years after HS entry	81820	0.611	0.488	100391	0.444	0.497
Any children 20 years after HS entry	81820	0.796	0.403	100391	0.683	0.465
# of children 5 years after HS entry	81820	0.020	0.151	100391	0.007	0.088
# of children 10 years after HS entry	81820	0.259	0.549	100391	0.127	0.392
# of children 15 years after HS entry	81820	1.015	0.965	100391	0.670	0.857
# of children 20 years after HS entry	81820	1.656	1.066	100391	1.317	1.080

Table 1: Descriptive Statistics – continued

Panel B: Student level background variables						
Mother has tertiary education	81820	0.3557	0.4787	100391	0.3911	0.488
Mother has upper secondary education	81820	0.3696	0.4827	100391	0.3648	0.4814
Mother has less than upper secondary education	81820	0.2558	0.4363	100391	0.2208	0.4148
Father has tertiary education	81820	0.3759	0.4844	100391	0.4305	0.4951
Father has upper secondary education	81820	0.3716	0.4832	100391	0.3526	0.4778
Father has less than upper secondary education	81820	0.2053	0.404	100391	0.1602	0.3668
First generation immigrant	81820	0.008	0.091	100391	0.012	0.107
Second generation immigrant	81820	0.006	0.075	100391	0.008	0.088
Child is adopted	81820	0.009	0.095	100391	0.007	0.084
Mother's age at birth	81052	26.429	4.769	99156	26.563	4.752
Mother <22 years at birth	81052	0.141	0.348	99156	0.133	0.340
Firstborn	81820	0.505	0.500	100391	0.516	0.500
Number of siblings	81689	1.472	0.908	100077	1.436	0.901
Lives with both parents at age 10	81820	0.861	0.346	100391	0.841	0.366
Panel C: School level variables						
Proportion female peers	81820	0.454	0.066	100391	0.447	0.067
Number of students in cohort	81820	107.878	30.776	100391	108.470	31.053
Cohort	81820	1987	4.338	100391	1987	4.345
Number of feeding municipalities	81820	6.330	3.304	100391	6.305	3.215
2+ high schools in municipality	81820	0.446	0.497	100391	0.459	0.498

NOTE: Annual labor earnings are measured in thousand Danish krones (1 USD = 6.6 DKK). Annual labor earnings are adjusted for 2015prices.

Table 2: Balancing Test I

Does Student Gender Predict Proportion of Female Peers?

	Dependent variable: Proportion female peers in high school cohort							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Female	0.0009	0.0014	0.0009	0.0012	0.0008	0.0012	0.0007	0.0011
	(0.0013)	(0.0010)	(0.0009)	(0.0009)	(0.0013)	(0.0010)	(0.0009)	(0.0009)
Observations	182211	182211	182211	182211	182211	182211	182211	182211
p-value (Female)	0.491	0.186	0.344	0.174	0.521	0.237	0.443	0.230
High school and cohort fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
School level controls	-	-	-	-	✓	✓	✓	✓
School-specific time trends	-	linear	quadratic	cubed	-	linear	quadratic	cubed

NOTE: The dependent variable in all columns is the proportion of female peers in the high school cohort of an individual. All columns include cohort fixed effects and school fixed effects. Following the Guryan, Kroft, and Notowidigdo (2009) correction method, we control for the leave-out mean of the proportion of female peers across cohorts within the school in all columns. *School level controls* included in Columns (5)-(8) include an indicator if any student in the cohort is older than 20 years at high school entry, dummies for number of students without information on gender (ranging from 0 to 2), indicators for large changes in cohort size compared to previous years, the female share in the language track, an indicator if the high school has no language track, indicators for exposure to experiment on course curriculum, and cubed cohort size.

Table 3: Balancing Test II

Does Proportion of Female Peers Predict Student Background Characteristics?

	(1)	(2)	(3)	(4)	(5)	(6) Across all models
Number of performed tests	38	38	38	38	38	190
Number significant at 1 percent level	1	1	0	0	0	2
Number significant at 5 percent level	1	1	0	1	1	4
Number significant at 10 percent level	4	5	2	4	2	17
Share significant at 1 percent	0.026	0.026	0.000	0.000	0.000	0.011
Share significant at 5 percent	0.026	0.026	0.000	0.026	0.026	0.021
Share significant at 10 percent	0.105	0.132	0.053	0.105	0.053	0.089
School level controls	-	✓	✓	✓	✓	
School-specific time trends	-	-	linear	quadratic	cubed	

NOTE: This Table is based on 190 separate OLS regressions shown in Appendix Table A1. All regressions include cohort fixed effects and school fixed effects. *School level controls* included in Columns (2)-(5) include an indicator if any student in the cohort is older than 20 years at high school entry, dummies for number of students without information on gender (ranging from 0 to 2), indicators for large changes in cohort size compared to previous years, the female share in the language track, an indicator if the high school has no language track, indicators for exposure to experiment on course curriculum, and cubed cohort size. Standard errors are clustered at the school-cohort level.

Table 4: Balancing Test III
School Level Autocorrelation in the Propotion Female Students

For what proportion of schools does the proportion of female students in t-1 significantly predict the proportion female students in t?

Proportion of school coefficients	(1)	(2)	(3)	(4)	Across all models
Significant at 1 percent level	1.57%	0.79%	0.79%	0.79%	0.98%
Significant at 5 percent level	2.36%	2.36%	3.15%	5.51%	3.35%
Significant at 10 percent level	6.30%	5.51%	8.66%	14.96%	8.86%
School-specific time trends	-	linear	quadratic	cubed	

NOTE: This table provides summary statistics of significance for 127 separate bivariate school-level regressions that only include the respective school-specific trend variable(s).

Table 5: The Impact of Peer Gender on STEM Enrollment and STEM Degree Completion

Panel A: STEM Enrollment	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female * Proportion female peers	-0.211***	-0.155***	-0.139***	-0.135***	-0.138***	-0.148***	-0.233***
	(0.026)	(0.026)	(0.026)	(0.026)	(0.026)	(0.027)	(0.036)
Male * Proportion female peers	0.056**	0.104***	0.086***	0.091***	0.092***	0.081***	
	(0.028)	(0.028)	(0.027)	(0.027)	(0.027)	(0.028)	
Female	-0.047***	-0.049***	-0.064***	-0.063***	-0.062***	-0.062***	-0.061***
	(0.017)	(0.017)	(0.016)	(0.016)	(0.016)	(0.016)	(0.016)
Observations	182211	182211	182211	182211	182211	182211	182211
Mean dependent variable women	0.210	0.210	0.210	0.210	0.210	0.210	0.210
Mean dependent variable men	0.378	0.378	0.378	0.378	0.378	0.378	0.378
p-values of test for gender equality of "Proportion female peers"	<.0001	<.0001	<.0001	<.0001	<.0001	<.0001	-
School fixed effects	-	✓	✓	✓	✓	✓	
Cohort fixed effects	-	✓	✓	✓	✓	✓	
Cohort controls	-	✓	✓	✓	✓	✓	
Individual & peer level controls	-	-	✓	✓	✓	✓	✓
School-specific time trends	-	-	-	linear	quadratic	cubed	-
Cohort-by-School fixed effects	-	-	-	-	-	-	✓

Panel B: STEM Completion	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female * Proportion female peers	-0.104***	-0.112***	-0.097***	-0.097***	-0.099***	-0.110***	-0.197***
	(0.021)	(0.022)	(0.022)	(0.022)	(0.022)	(0.023)	(0.030)
Male * Proportion female peers	0.122***	0.106***	0.093***	0.092***	0.093***	0.082***	
	(0.025)	(0.024)	(0.023)	(0.023)	(0.023)	(0.024)	
Female	-0.016	-0.019	-0.032**	-0.033**	-0.031**	-0.031**	-0.030**
	(0.014)	(0.014)	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)
Observations	182211	182211	182211	182211	182211	182211	182211
Mean dependent variable women	0.135	0.135	0.135	0.135	0.135	0.135	0.135
Mean dependent variable men	0.253	0.253	0.253	0.253	0.253	0.253	0.253
p-values of test for gender equality of "Proportion female peers"	<.0001	<.0001	<.0001	<.0001	<.0001	<.0001	-
School fixed effects	-	✓	✓	✓	✓	✓	
Cohort fixed effects	-	✓	✓	✓	✓	✓	
Cohort controls	-	✓	✓	✓	✓	✓	
Individual & peer level controls	-	-	✓	✓	✓	✓	✓
School-specific time trends	-	-	-	linear	quadratic	cubed	-
Cohort-by-School fixed effects	-	-	-	-	-	-	✓

NOTE: The dependent variable in all columns of Panel A is an indicator for whether the student ever enrolled in a STEM program in college within 20 years after high school entry. The dependent variable in all columns of Panel B is an indicator for whether the student's highest completed education is at least at the college level and is within STEM. Colum (7) does not include peer-level variables because these are highly collinear with the cohort-by-high school fixed effects. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Impact of High School Gender Composition on Graduation in Various Fields

Panel A: STEM Subfields	Biology	Math / Physics	ICT / Engineering	Manufacturing / Construction
	(1)	(2)	(3)	(4)
Female * Proportion female peers	-0.030***	-0.026**	-0.019	-0.023**
	(0.010)	(0.010)	(0.015)	(0.010)
Male * Proportion female peers	0.011	0.030***	0.039**	0.012
	(0.007)	(0.011)	(0.019)	(0.011)
Female	0.039***	0.012**	-0.083***	-0.000
	(0.005)	(0.006)	(0.010)	(0.006)
Observations	182211	182211	182211	182211
Mean dependent variable women	0.035	0.027	0.044	0.029
Mean dependent variable men	0.015	0.040	0.153	0.045
p-values of test for gender equality of "Proportion female peers"	0.001	<.0001	0.008	0.013

Panel B: Other Fields of Study	Health Sciences	Education	Arts / Humanities	Social Sciences	Business / Law
	(1)	(2)	(3)	(4)	(5)
Female * Proportion female peers	0.110***	0.022	0.001	-0.019	0.017
	(0.024)	(0.015)	(0.014)	(0.013)	(0.019)
Male * Proportion female peers	-0.077***	-0.037***	0.014	0.004	0.004
	(0.015)	(0.012)	(0.011)	(0.013)	(0.019)
Female	0.085***	0.007	0.020***	-0.001	-0.045***
	(0.012)	(0.008)	(0.007)	(0.007)	(0.011)
Observations	182211	182211	182211	182211	182211
Mean dependent variable women	0.217	0.075	0.058	0.049	0.093
Mean dependent variable men	0.047	0.040	0.045	0.062	0.134
p-values of test for gender equality of "Proportion female peers"	<.0001	0.001	0.407	0.144	0.576

NOTE: All models control for linear school-specific time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. All results are robust to the inclusion of cohort-by-school fixed effects. Standard errors clustered at the school-cohort level are in parentheses. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table 7: Mechanisms and Heterogeneity

Subgroup:	Full sample		No parent has college degree	1+ parent has college degree	Father has STEM education	Mother has STEM education
Dependent variable:	Std. High School GPA	STEM completion	STEM completion	STEM completion	STEM completion	STEM completion
	(1)	(2)	(3)	(4)	(5)	(6)
Female * Proportion female peers	-0.006	-0.097***	-0.128***	-0.067**	-0.112***	0.072
	(0.062)	(0.022)	(0.029)	(0.031)	(0.038)	(0.100)
Male * Proportion female peers	0.126**	0.092***	0.088***	0.094***	0.052	0.238**
	(0.058)	(0.023)	(0.032)	(0.030)	(0.040)	(0.101)
Female	0.061	-0.033**	-0.020	-0.047***	-0.064***	-0.031
	(0.038)	(0.013)	(0.019)	(0.018)	(0.023)	(0.057)
Observations	159603	182211	83783	98428	60544	9955
Mean dependent variable women	-0.011	0.135	0.165	0.103	0.158	0.207
Mean dependent variable men	0.007	0.253	0.280	0.217	0.298	0.307
p-values of test for gender equality of "Proportion female peers"	0.110	<.0001	0.008	0.001	0.001	0.192

NOTE: The dependent variable in column (1) is the grade point average (GPA) at the end of high school, standardized with a mean of zero and standard deviation of one, and is observed for those students who completed general academic high school. All models control for school-specific linear time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years, as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. All results are robust to the inclusion of cohort-by-school fixed effects. Standard errors clustered at the school-cohort level are in parentheses. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table 8: Impact of High School Gender Composition on Labor Market Outcomes

Panel A: Cohort and School fixed effects	Working in STEM occupation		Earnings percentile	
	11–15 years after HS entry (1)	16–20 years after HS entry (2)	11–15 years after HS entry (3)	16–20 years after HS entry (4)
Female * Proportion female peers	-0.032*	-0.050**	-3.898***	-6.655***
	(0.019)	(0.020)	(1.362)	(1.471)
Male * Proportion female peers	0.020	0.030	-1.222	0.082
	(0.022)	(0.022)	(1.431)	(1.433)
Female	-0.124***	-0.117***	-9.725***	-11.538***
	(0.012)	(0.013)	(0.855)	(0.911)
Observations	181022	178240	181012	178229
Mean dependent variable women	0.103	0.109	52.178	56.633
Mean dependent variable men	0.246	0.255	62.567	70.615
p-values of test for gender equality of "Proportion female peers"	0.048	0.004	0.148	0.001

Panel B: Cohort-by-School fixed effects	Working in STEM occupation		Earnings percentile	
	11–15 years after HS entry (1)	16–20 years after HS entry (2)	11–15 years after HS entry (3)	16–20 years after HS entry (4)
Female * Proportion female peers	-0.059**	-0.088***	-3.239*	-7.508***
	(0.027)	(0.029)	(1.892)	(2.032)
Female	-0.122***	-0.113***	-9.462***	-11.196***
	(0.012)	(0.013)	(0.864)	(0.917)
Observations	181022	178240	181012	178229
Mean dependent variable women	0.103	0.109	52.178	56.633
Mean dependent variable men	0.246	0.255	62.567	70.615

NOTE: The dependent variable in columns (1) and (2) is an indicator for working in a STEM occupation; it takes the value one if the individual for at least half the period works in a STEM occupation within the Danish version of ISCO codes 21, 25, 31 or 35 (Science and Engineering Professionals, Information and Communications Technology Professionals, Science and Engineering Associate Professionals, or Information and Communications Technicians). The dependent variable in columns (3) and (4) is the average of the individual's labor earnings percentile during the five-year period, calculated by year of birth and age using the entire Danish population as a reference group. All models in Panel A control for school-specific linear time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years, as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. All models in Panel B control for cohort-by-school fixed effects and individual controls. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 9: Impact of High School Gender Composition on Fertility

Panel A: Cohort and School fixed effects	Dependent variable: Any children				Dependent variable: Number of children			
	5 years after HS entry	10 years after HS entry	15 years after HS entry	20 years after HS entry	5 years after HS entry	10 years after HS entry	15 years after HS entry	20 years after HS entry
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Female * Proportion female peers	0.009	0.065***	0.023	0.013	0.011	0.103***	0.210***	0.168***
	-0.008	-0.022	-0.029	-0.025	-0.008	-0.03	-0.053	-0.062
Male * Proportion female peers	-0.009*	-0.065***	-0.085***	-0.051**	-0.007	-0.079***	-0.177***	-0.152**
	-0.005	-0.018	-0.027	-0.025	-0.005	-0.023	-0.048	-0.06
Female	0.004	0.040***	0.114***	0.081***	0.005	0.047***	0.161***	0.181***
	-0.004	-0.012	-0.017	-0.015	-0.004	-0.016	-0.03	-0.035
Observations	182211	182211	182211	182211	182211	182211	182211	182211
Mean dependent variable women	0.019	0.208	0.611	0.796	0.02	0.259	1.015	1.656
Mean dependent variable men	0.007	0.108	0.444	0.683	0.007	0.127	0.67	1.317
p-values of test for gender equality of "Proportion female peers"	0.027	<.0001	0.003	0.044	0.042	<.0001	<.0001	<.0001

Panel B: Cohort-by-School fixed effects	Dependent variable: Any children				Dependent variable: Number of children			
	5 years after HS entry	10 years after HS entry	15 years after HS entry	20 years after HS entry	5 years after HS entry	10 years after HS entry	15 years after HS entry	20 years after HS entry
Female * Proportion female peers	0.017***	0.126***	0.096***	0.056*	0.017*	0.181***	0.366***	0.294***
	-0.008	-0.027	-0.037	-0.032	-0.009	-0.035	-0.067	-0.079
Female	0.004	0.042***	0.120***	0.085***	0.005	0.048***	0.172***	0.194***
	-0.004	-0.012	-0.017	-0.015	-0.004	-0.016	-0.03	-0.035
Observations	182211	182211	182211	182211	182211	182211	182211	182211
Mean dependent variable women	0.019	0.208	0.611	0.796	0.02	0.259	1.015	1.656
Mean dependent variable men	0.007	0.108	0.444	0.683	0.007	0.127	0.67	1.317

NOTE: Five years after high school entry, individuals should be at the beginning of their college studies. Ten years after high school entry, individuals are around 26 years old and should have completed college education. Fifteen years after high school, individuals have been in the labor market for approximately 5 years if they attended college. Twenty years after high school, individuals are about 36 years old. All models in Panel A control for school-specific time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. All models in Panel B include cohort-by-school fixed effects and individual controls. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 10: Impact of High School Peers on Earnings—Including Fertility and STEM Controls

Panel A: Cohort and School fixed effects — Dependent Variable: Earnings Percentile

	Main Result (1)	Fertility Controls (2)	STEM Controls (3)	Fertility and STEM Controls (4)
Female * Proportion female peers	-6.655***	-4.001***	-3.624***	-1.370
	(1.471)	(1.423)	(1.325)	(1.280)
Male * Proportion female peers	0.082	-1.022	-0.185	-1.206
	(1.433)	(1.371)	(1.299)	(1.256)
Female	-11.538***	-1.575*	-13.494***	-3.943***
	(0.911)	(0.900)	(0.817)	(0.806)
Observations	178229	178229	178229	178229
Mean dependent variable women	56.633	56.633	56.633	56.633
Mean dependent variable men	70.615	70.615	70.615	70.615
p-values of test for gender equality of "Proportion female peers"	0.001	0.119	0.054	0.924

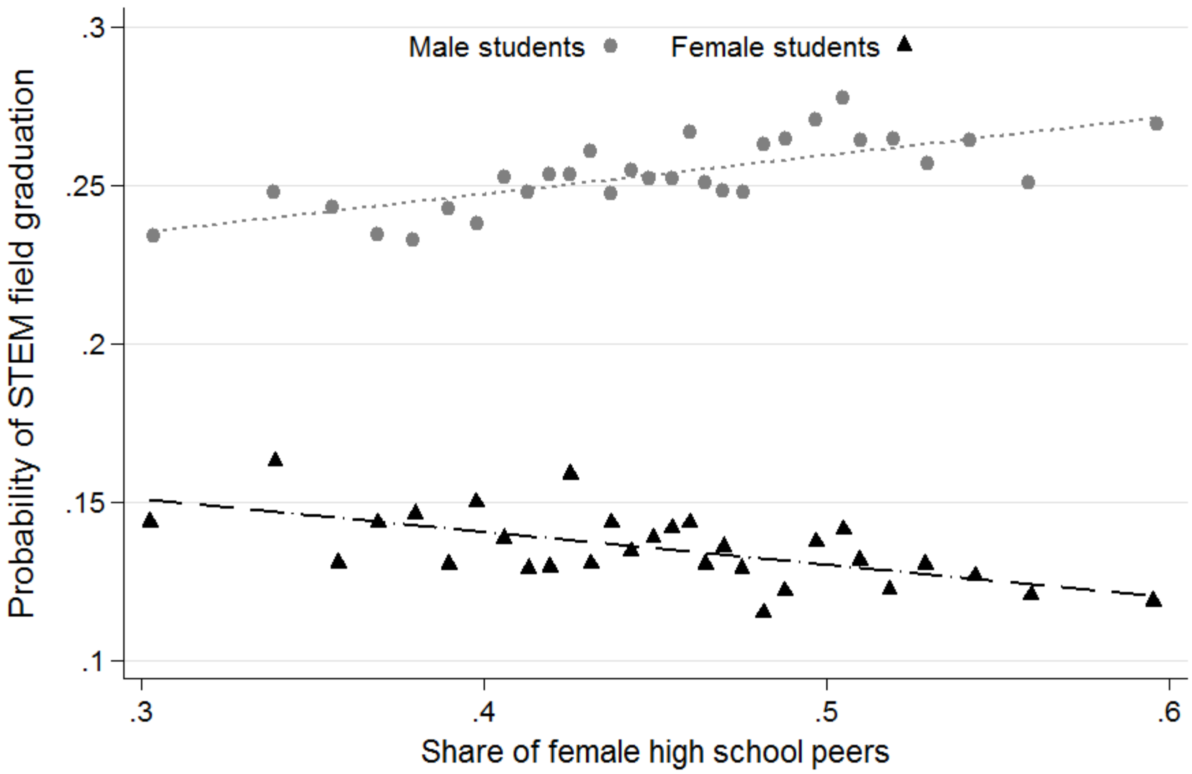
Panel B: Cohort-by-School fixed effects — Dependent Variable: Earnings Percentile

	Main Result (1)	Fertility Controls (2)	STEM Controls (3)	Fertility and STEM Controls (4)
Female * Proportion female peers	-7.508***	-3.646*	-3.925**	-0.590
	(2.032)	(1.942)	(1.815)	(1.746)
Female	-11.196***	-1.306	-13.282***	-3.773***
	(0.917)	(0.904)	(0.821)	(0.808)
Observations	178229	178229	178229	178229
Mean dependent variable women	56.633	56.633	56.633	56.633
Mean dependent variable men	70.615	70.615	70.615	70.615

NOTE: The dependent variable in all columns is the average labor earnings percentile 16–20 years after high school entry, calculated by year of birth and age using the entire Danish population as a reference group. All models in Panel A control for school-specific linear time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. All models in Panel B control for cohort-by-school fixed effects and individual controls. *Fertility Controls* include gender-specific dummies for number of children 15 and 20 years after high school entry. *STEM Controls* include gender-specific dummies for working in STEM occupations and having completed a STEM college degree and their interactions. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

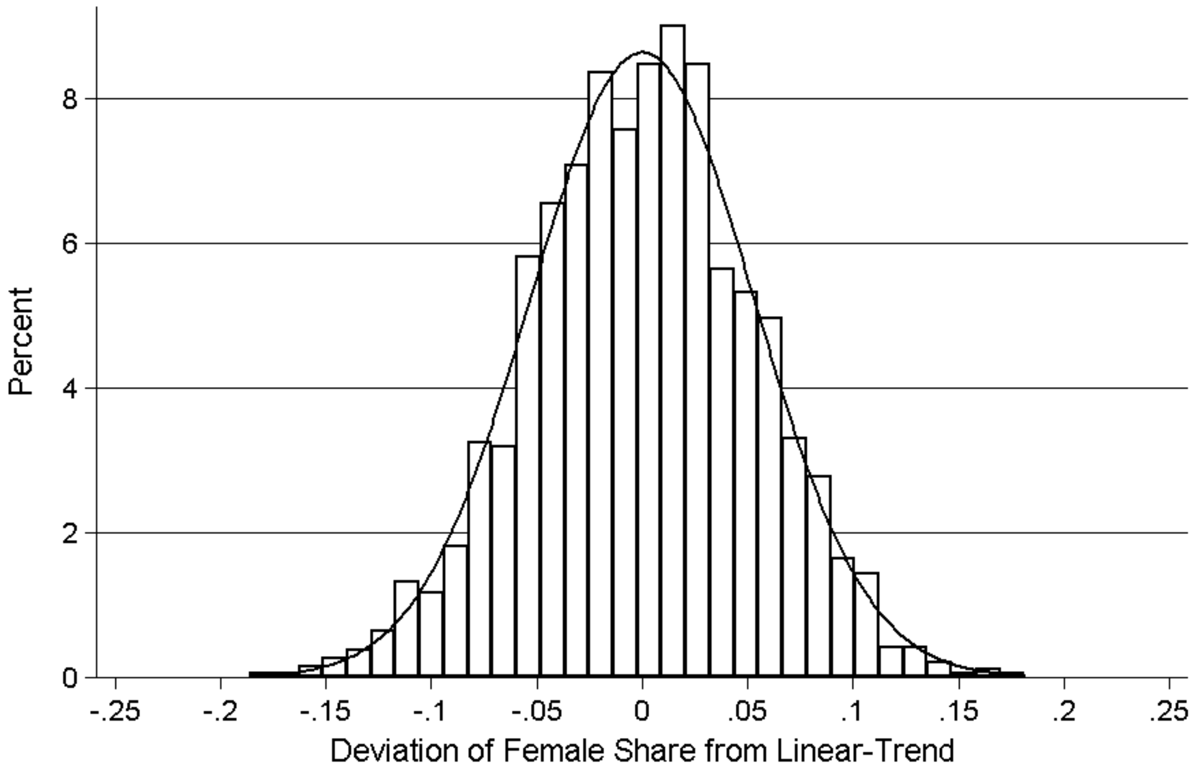
FIGURES

Figure 1: Correlation between Proportion of Female Peers and STEM Degree Completion



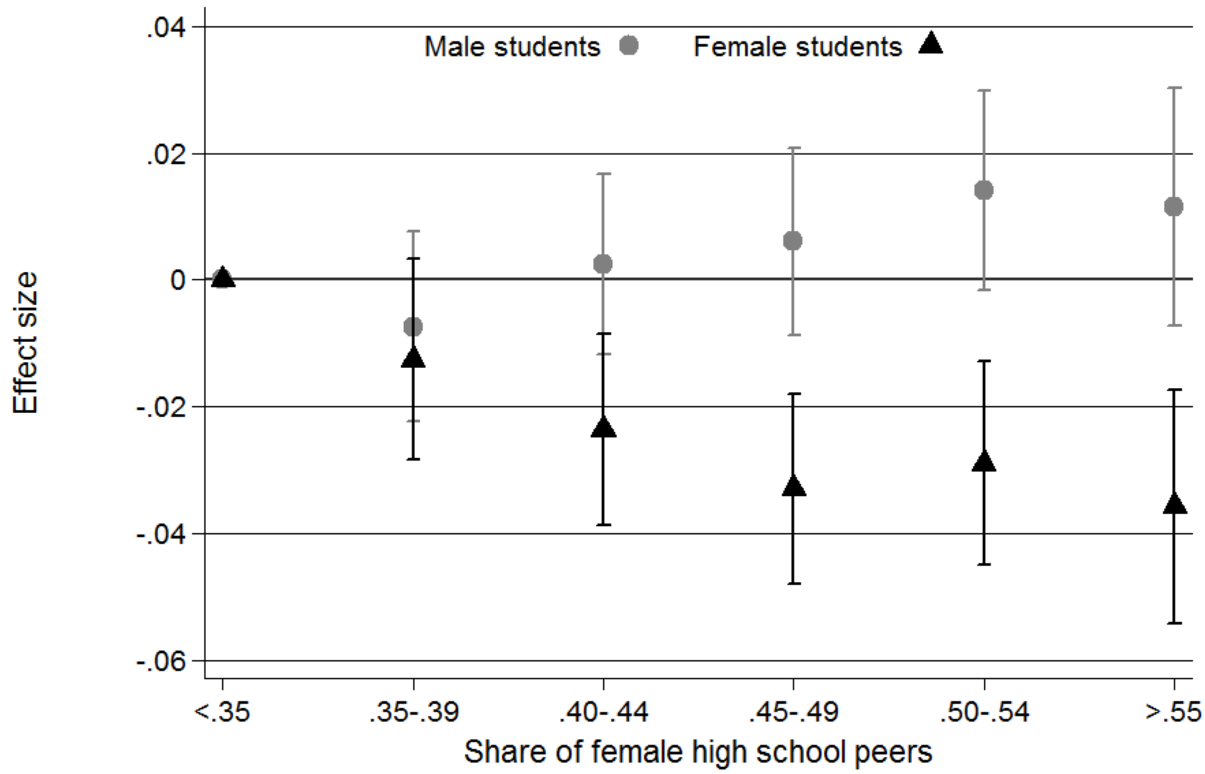
NOTE: The graph shows a bin scatter plot by gender using 30 bins. STEM Degree is measured as highest completed degree at the college level or higher 20 years after high school entry.

Figure 2: Year-to-Year Variation in the Proportion of Female High School Peers Within High Schools



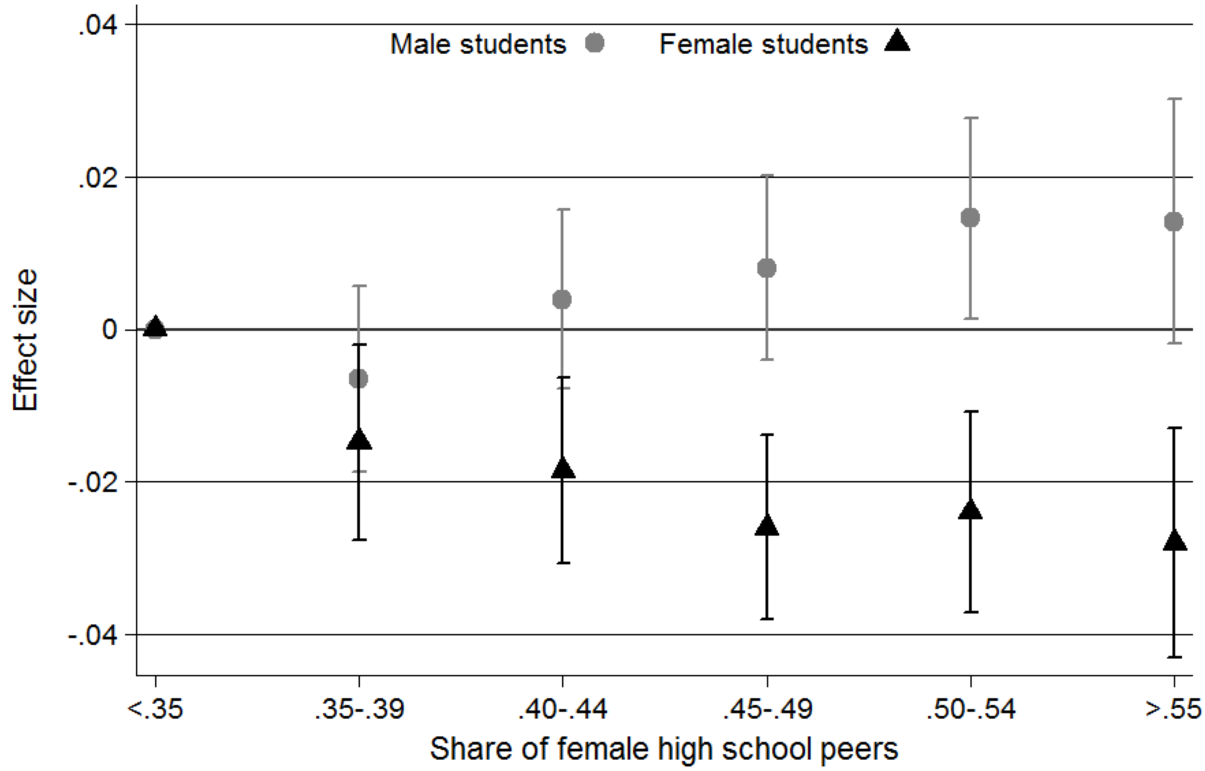
NOTE: This figure illustrates the year-to-year variable in the proportion of female high school peers within high schools, plotted together with the normal distribution. More precisely, it plots the predicted proportion of female peers at the school-cohort level from a regression regressing the proportion of female peers on cohort and school fixed effects and school-specific linear time trends. Each high school-cohort represents one observation.

Figure 3: High School Gender Composition and STEM Enrollment – Non-linear Effects



NOTE: This figure shows point estimates obtained from OLS regression. Instead of including the continuous measure of the proportion female high school peers interacted with the gender dummy, this regression includes five dummies for the high school gender composition interacted with gender. The dependent variable is STEM study enrollment. Vertical lines refer to the 95 percent confidence intervals. The model controls for school-specific linear time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years, as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1.

Figure 4: High School Gender Composition and STEM Graduation – Non-linear Effects



NOTE: This figure shows point estimates obtained from OLS regression. Instead of including the continuous measure of the proportion female high school peers interacted with the gender dummy, this regression includes five dummies for the high school gender composition interacted with gender. The dependent variable is STEM graduation. Vertical lines refer to the 95 percent confidence intervals. The model controls for school-specific linear time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years, as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1.

APPENDIX

Table A1: Complete Balancing Test Including All Individual Level Variables

Based on 190 separate regressions	(1)	(2)	(3)	(4)	(5)
Age at high school entry	-0.001	-0.002	0.002	0.008	0.018
	(0.032)	(0.032)	(0.029)	(0.029)	(0.029)
Mother has less than upper-secondary education	0.024	0.024	0.004	0.004	-0.011
	(0.020)	(0.020)	(0.018)	(0.019)	(0.019)
Mother has upper-secondary education	-0.005	-0.005	-0.006	0.001	0.003
	(0.023)	(0.023)	(0.021)	(0.022)	(0.022)
Mother has tertiary education	-0.020	-0.019	0.005	-0.002	0.013
	(0.021)	(0.021)	(0.021)	(0.021)	(0.021)
Mother education unknown	0.001	0.000	-0.003	-0.003	-0.006
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Father has less than upper-secondary education	0.019	0.017	-0.008	-0.013	-0.015
	(0.017)	(0.017)	(0.016)	(0.016)	(0.017)
Father has upper-secondary education	0.014	0.020	0.016	0.020	0.019
	(0.023)	(0.023)	(0.022)	(0.022)	(0.023)
Father has tertiary education	-0.018	-0.021	0.009	0.013	0.014
	(0.022)	(0.022)	(0.021)	(0.021)	(0.022)
Father education unknown	-0.015	-0.015	-0.017*	-0.019*	-0.018*
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
Mother Education field	0.005	0.005	0.013	0.014	0.018
	(0.012)	(0.012)	(0.012)	(0.012)	(0.013)
Mother Humanities field	-0.015*	-0.015*	-0.011	-0.014*	-0.010
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
Mother Social Sciences field	0.001	-0.000	0.004	0.005	0.006
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Mother Business, Admin, and Law field	-0.033*	-0.033*	-0.028	-0.026	-0.020
	(0.019)	(0.019)	(0.018)	(0.018)	(0.018)
Mother STEM field	-0.003	-0.002	0.002	0.003	0.002
	(0.009)	(0.009)	(0.009)	(0.009)	(0.010)
Mother Life Sciences field	0.000	0.000	0.000	0.000	0.001
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Mother Health and Welfare field	0.008	0.011	0.012	0.008	0.007
	(0.019)	(0.019)	(0.019)	(0.019)	(0.019)
Mother Service field	0.001	0.001	-0.000	0.001	0.002
	(0.006)	(0.006)	(0.006)	(0.006)	(0.006)
Mother No field	0.034*	0.033	0.012	0.013	0.001
	(0.020)	(0.020)	(0.019)	(0.019)	(0.019)
Individual level and high school level controls	-	✓	✓	✓	✓
School-specific time trends	-	-	linear	quadratic	cubed

Table A1— continued

	(1)	(2)	(3)	(4)	(5)
Father Education field	0.001	0.001	0.007	0.009	0.007
	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Father Humanities field	-0.019***	-0.020***	-0.010	-0.009	-0.008
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Father Social Sciences field	-0.008	-0.009*	-0.008	-0.009*	-0.008
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Father Business, Admin, and Law field	0.010	0.012	0.022	0.026*	0.023
	(0.015)	(0.015)	(0.015)	(0.015)	(0.016)
Father STEM field	0.016	0.017	0.011	0.005	0.008
	(0.021)	(0.021)	(0.020)	(0.020)	(0.021)
Father Life Sciences field	0.004	0.005	0.007	0.014**	0.013*
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Father Health and Welfare field	-0.008	-0.008	-0.004	-0.004	-0.002
	(0.010)	(0.010)	(0.010)	(0.010)	(0.010)
Father Service field	0.005	0.005	0.004	0.001	-0.003
	(0.009)	(0.009)	(0.009)	(0.009)	(0.009)
Father No field	0.015	0.015	-0.014	-0.017	-0.017
	(0.017)	(0.018)	(0.016)	(0.017)	(0.017)
Child is adopted	0.004	0.004	0.003	0.003	0.003
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Lives with both parents at age 10	0.006	0.010	0.010	0.006	0.005
	(0.015)	(0.015)	(0.015)	(0.015)	(0.014)
First-generation immigrant	0.001	0.000	-0.004	-0.001	-0.001
	(0.005)	(0.005)	(0.005)	(0.004)	(0.005)
Second-generation immigrant	0.004	0.003	-0.004	-0.003	-0.004
	(0.005)	(0.005)	(0.004)	(0.004)	(0.004)
Firstborn	-0.039*	-0.041*	-0.038*	-0.024	-0.009
	(0.022)	(0.022)	(0.022)	(0.022)	(0.023)
Number of siblings	0.007	-0.004	-0.028	-0.038	-0.059
	(0.039)	(0.039)	(0.038)	(0.039)	(0.041)
Number of siblings squared	-0.029	-0.100	-0.345	-0.402	-0.569**
	(0.267)	(0.266)	(0.257)	(0.265)	(0.282)
Mother's age at birth	0.267	0.308	0.274	0.063	-0.035
	(0.265)	(0.265)	(0.250)	(0.248)	(0.253)
Mother's age at birth squared	15.921	17.982	14.983	7.406	1.763
	(13.486)	(13.493)	(12.545)	(12.380)	(12.625)
Mother <22 years at birth	-0.013	-0.015	-0.011	-0.001	0.006
	(0.016)	(0.016)	(0.016)	(0.016)	(0.016)
Mother's age unknown	0.003	0.003	0.001	0.005	0.004
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
School level controls	-	✓	✓	✓	✓
School-specific time trends	-	-	linear	quadratic	cubed

NOTE: Each cell in this table is estimated with a separate regression including including school and cohort fixed effects. The dependent variable in each cell is the proportion of female high school peers. *School level controls* included in Columns (2)-(5) are an indicator if any student in the cohort is older than 20 years at high school entry, dummies for number of students without information on gender (ranging from 0 to 2), indicators for large changes in cohort size compared to previous years, the female share in the language track, an indicator if the high school has no language track, indicators for exposure to experiment on course curriculum, and cubed cohort size. Standard errors in parentheses are clustered at the school-cohort level. * p < 0.1, ** p < 0.05, *** p < 0.01.

Table A2: The Impact of Peer Gender on High School Graduation, College Enrollment, and Higher Education Degree Completion

	(1) Completed academic high school	(2) Ever enrolled in higher education	(3) Completed higher education degree
Female * Proportion female peers	0.017	-0.010	0.000
	(0.017)	(0.024)	(0.027)
Male * Proportion female peers	0.017	-0.030	-0.035
	(0.017)	(0.022)	(0.026)
Female	0.024**	0.005	0.045***
	(0.010)	(0.014)	(0.016)
N	182211	182211	182211
Mean	0.909	0.789	0.677
p-values of test for gender equality of "Proportion female peers"	0.992	0.523	0.315

NOTE: The dependent variable in Column (1) is equal to one if the student completed academic high school within 5 years after high school entry. The dependent variable in Column (2) is equal to one if the student ever enrolled in college studies and the dependent variable in Column (3) is equal to one if the student ever completed any college education. All models control for school-specific time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A3: Robustness Check—Main Results by Number of Schools in the Municipality

Subgroup	Full sample	Municipality with only 1 high schools	Municipality with 2+ high schools
Dependent variable:	STEM completion	STEM completion	STEM completion
	(1)	(2)	(3)
Female * Proportion female peers	-0.097***	-0.088***	-0.097***
	(0.022)	(0.030)	(0.032)
Male * Proportion female peers	0.092***	0.107***	0.064**
	(0.023)	(0.032)	(0.033)
Female	-0.033**	-0.036*	-0.037**
	(0.013)	(0.019)	(0.019)
Observations	182211	99599	82612
Mean	0.200	0.200	0.199
p-values of test for gender equality of "Proportion female peers"	<.0001	0.001	0.046

NOTE: All models control for school-specific time trends, cohort fixed effects, school fixed effects, cubed cohort size, indicators for large cohort size changes compared to previous years as well as a large set of individual and leave-out-mean peer controls shown in Panel B of Table 1. Standard errors clustered at the school-cohort level are in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Chapter 3 – Gender Gaps in the Effects of Childhood Family Environment: Do They Persist into Adulthood?

Gender Gaps in the Effects of Childhood Family Environment: Do They Persist into Adulthood?

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Abstract

We examine the differential effects of family disadvantage on the education and adult labor market outcomes of men and women using high-quality administrative data on the entire population of Denmark born between 1966 and 1995. We link parental education and family structure during childhood to male-female and brother-sister differences in adolescent outcomes, educational attainment, and adult earnings and employment. Our results are consistent with U.S. findings that boys benefit more from an advantageous family environment than do girls in terms of grade-school outcomes. Father's education, which has not been examined in previous studies, is particularly important for sons. However, we find a very different pattern of parental influence on adult outcomes. Gender gaps in educational attainment, employment, and earnings are increasing in maternal education, benefiting daughters. Paternal education decreases the gender gaps in educational attainment (favoring sons) and labor market outcomes (favoring daughters). We conclude that differences in the behavior of school-aged boys and girls may be poor proxies for differences in skills that drive longer-term outcomes.

JEL classification: I20, J1, J2, J3

Keywords: Gender gap, parental education, family structure, education, labor market outcomes.

1 Introduction

Over the past century, barriers to women's educational and employment opportunities have been dramatically lowered in most of the developed world. Women continue to have lower rates of labor force participation and earn lower pay than men, but new gender gaps that favor women have opened up in education. Young men lag behind young women in academic achievement and contributing factors include less engagement in school, a gap in homework hours and the substitution of time spent playing video games for time spent reading (OECD, 2015). Women are now more likely than men to complete secondary education and to graduate from college in almost all OECD countries. In the United States, 39 percent of women aged 25 to 29 have a Bachelor's degree or more, compared to 32 percent of men (U.S. Census Bureau, 2015).

Recent studies have focused on the behavioral differences between school-aged boys and girls, arguing that a gender gap in "non-cognitive skills" contributes to the scholastic underperformance of boys by increasing the costs of school persistence and performance (Goldin et al., 2006; Becker et al., 2010). Family disadvantage is strongly negatively associated with early social and behavioral skills for both boys and girls, and it has been suggested that trends in family structure, and in particular the increasing prevalence of single parent families, may have a particularly deleterious effect on the skill development of boys (Bertrand and Pan, 2013; Autor and Wasserman, 2013).

Autor et al. (2016) examine this hypothesis using sibling fixed-effects models and a sample of students in Florida, and find that early family structure and mother's education do have significantly larger effects on a variety of school outcomes for boys than for their sisters. However, there is also evidence that the greater impact of family background on boys is most relevant for school-age behavior in the United States, and does not extend to longer-term outcomes such as educational attainment (Lundberg, 2016). With our analysis, we contribute to this literature in three important ways. First, we re-examine and confirm gender differences in the impacts of family environment on school-age outcomes for Denmark, another OECD country with different social institutions and lower poverty prevalence, especially among single-parent families. Second, our main contribution is to examine a broad range of adult outcomes for the total population as well as for large samples of full siblings. Third, the richness of the data makes it possible to study potential differences in family environment effects across cohorts. Administrative data on the entire population of Denmark from 1980 to 2015 with cohorts born from 1966 to 1995 enables us to link parental education and family structure during childhood to male-female differences in adolescent outcomes, educational attainment, and adult earnings and employment. A significant advantage of the Danish administrative data is that we are able to add paternal education, which is not available for large subsets of the American samples, to our indicators of family background.

We find, as do Autor et al. (2016) and Lundberg (2016), that adolescent boys appear to be more sensitive than girls to family environment. However, we find a very different pattern of parental influence on adult outcomes such as educational attainment, college graduation, employment, and earnings. Maternal education consistently has a greater impact on the education and employment of daughters relative to sons and this effect is stable across cohorts. Paternal education has some significant, though smaller, effects on the gender education gap that favor sons (and that decline over time). These positive effects of same-sex parental education may reflect role-modelling. Father's education has larger positive effects, however, on the employment and earnings of daughters, which may indicate that female labor market behavior in Denmark is more elastic than men's with respect to early influences. The effects of family structure on adult outcomes vary, and are both small and less consistent across samples than the impacts of parental education. Having married parents at birth tends to increase the relative education and earnings of men, but to reduce their relative probability of college graduation.

Estimates based on the total population are similar to those obtained from a sample of full siblings controlling for family fixed effects. This suggests that the selection of boys and girls across different family types is not biasing our estimates of the gender gap in the effects of family environment in the full sample.

We conclude that, although there are gender differences in responses to parental resources and family structure, they do not conform to the simple story that the skill development of boys is particularly vulnerable to family disadvantage. Our results are consistent with an alternative hypothesis in which maternal education and other family resources have a moderating effect on the outcomes of behavioral and developmental problems in school that are much more typical of boys than girls. These parental influences become less important as the children become adults, and we find little support for the hypothesis that these early behavior gaps imply less long-term skill acquisition by boys, relative to girls.

2 Family Background and Child Outcomes: Is There a Gender Dimension?

Boys begin school with less-developed social and behavioral skills than girls, and these gaps persist through elementary school and explain much of the gender differential in early academic outcomes (DiPrete and Jennings, 2012). Girls consistently receive higher grades, are less likely to repeat grades or to be placed in special education classes, and are less likely to get in trouble at school. There are clear behavioral patterns underlying these disparate outcomes—girls spend more time on homework, are more likely to read for pleasure, and exhibit a greater degree of self-discipline in school.[1] Attempts to explain the emergence of a gender gap favoring women in college attendance and completion have appealed to these gender differences in academic achievement and school discipline as evidence of a "non-cognitive skill" deficit that increases the effective costs of attending and succeeding in school for boys (Goldin et al.,

[1] Duckworth and Seligman (2006) use several measures of self-discipline to document this gender difference, including self-reports, teacher and parent reports, and a delay of gratification test.

2006; Becker et al., 2010).[2]

In addition to this gender skill gap, there are also strong socioeconomic gradients in early social skills, attention, and school engagement. These skill differences can explain a portion of the socioeconomic differences in young adult outcomes such as arrests and high school completion (Duncan and Magnuson, 2011). Autor and Wasserman (2013) suggest a new explanation for the trend in the relative educational attainment of men and women based on these socioeconomic skill differentials and trends in family structure. They hypothesize that, as the prevalence of single parent families has increased in the U.S. (and elsewhere), economic stresses have increased for children in lower income households and their access to paternal time and attention has decreased. If the skill development of boys is affected more by father absence or family disadvantage than the skill development of girls, then changes in the living arrangements of children over time may play a role in the growing education gender gap.[3] Bertrand and Pan (2013) provide supportive empirical evidence, showing that living with a single mother or a young mother has a much larger effect on externalizing behavior and school suspensions for boys than for girls. They interpret the negative behavioral impact of father absence and young mothers as evidence that the non-cognitive skills development of boys is particularly sensitive to family disadvantage.

Autor et al. (2016) re-examine this "vulnerable boys" hypothesis using data for a large sample of children in Florida that links birth certificates with academic and health records. Using a variety of measures of family environment (including mother's education, marital status at birth, father presence, and an SES index), neighborhood

[2]Other studies have investigated possible gender gaps in the benefits of education. Becker et al. (2010) conclude that the monetary returns to education are still lower for women than for men. However, (Browning et al., 2014) suggest that women may benefit more from education through returns in the marriage market, or through a gender wage gap that declines in education. In the latter case, women may invest more in formal schooling to escape from labor market discrimination.

[3]It is not clear what the mechanisms might be that make boys more vulnerable to adverse environments in childhood. One possibility is that gender differences in developmental trajectories may make girls, who enter school more mature in language skills and emotional regulation, inherently more resilient to disadvantage. Alternatively, there may be socioeconomic differences parental investment strategies that lead low-income parents to favor girls. Bertrand and Pan (2013) find that single mothers spend more time with daughters than with sons and report less emotional closeness with sons. Finally, there may be cultural factors that lead boys, in particular, to develop negative attitudes to school in low income or single parent families or that inhibit the educational aspirations of boys relative to girls (DiPrete and Buchmann, 2013).

income and school quality, they find that early family structure and mother's education do have significantly larger effects on a variety of school outcomes for boys than for their sisters, including school suspensions and absences, in both OLS and family fixed effects models. There is a larger payoff for boys to having a college graduate mother for a broad set of academic outcomes, including kindergarten readiness and grades. They find similar patterns of differential gender impacts of low-income neighborhoods and poor-quality schools, and conclude that family disadvantage has larger impacts on the outcomes of boys relative to girls throughout school. Though they are unable to examine later outcomes, including college attainment, earnings, and labor force participation, Autor et al. (2016) suggest that early gender differences in behavioral and school outcomes are likely to have implications for adult outcomes.

Other studies cast some doubt on this final speculation, however. Riphahn and Schwientek (2015) find no link between family background and gender differences in educational attainment in German micro-data. Lundberg (2016), using the National Longitudinal Study of Adolescent to Adult Health (Add Health), finds that differential vulnerability to father absence among school-aged boys and girls depends upon the outcome: boys appear more responsive to father absence in externalizing behaviors (problems in school, school suspensions) while girls appear more vulnerable when outcomes are related to internalizing behavior (depression). However, neither of these patterns of adolescent response have any significant implications for educational outcomes: father absence has no differential impact on college graduation rates of men and women in cross-sectional or sibling fixed-effects models.[4] The question of whether gender differences in the effects of childhood environment persist into adulthood has potentially important implications for early childhood and school interventions, which may be designed with a focus on boys or girls.

Using the large samples available in Danish administrative data, we test the hy-

[4]Fan et al. (2015) take a different approach to the emerging gender gap, postulating that boys may be more adversely affected by mother's employment in childhood. They find evidence for a more positive association between mother's work and girl's education in Norwegian administrative data using family fixed-effect models. They do not, however, control for mother's education, which we find is a stronger predictor of daughters' outcomes than of sons'.

pothesis that males benefit more from mother's and father's education and from having married parents at birth than do females in terms of a broad range of adult outcomes. Denmark has experienced trends in relative male and female educational attainment and single-parent households that are similar to those in the U.S., though the educational and labor market environments are distinct.[5] The more comprehensive social safety net and the well-defined obligations of non-resident parents may moderate the impacts of family disadvantage on child outcomes.[6] Though female labor force participation rates in Denmark are high, women are more likely to work part-time than in the U.S. and also more likely to work in the public sector. On the other hand, we expect that any developmental process that renders boys more vulnerable to adverse family environments should be a very general one that is manifest in diverse institutional environments, though effect sizes are likely to differ. Landersø and Heckman (2016), for example, find that despite social policy differences, the influence of family background on educational attainment is similar in Denmark and the U.S.

3 Data

We use Danish administrative data covering the entire population born in Denmark between 1966 and 1995 to examine both outcomes during adolescence and the longer-term consequences of parental resources and family structure in early life. One important feature of this dataset is that we are able to link each child to his or her biological parents (both mother and father) and siblings. Moreover, we observe educational and labor market outcomes for each year, and can track with whom each individual lives.

[5]Appendix Figure A1 shows the growth in non-traditional family structures experienced by Danish children at age 12 and Appendix Figure A2 illustrates the reversal in the gender gap in highest completed education by age 31 for Denmark.

[6]See e.g. Rossin-Slater and Wüst (2014) on child support obligations in Denmark.

3.1 Family Childhood Environment

We measure three dimensions of childhood family environment: mother's education, father's education, and marital status at birth. In the administrative data, we observe the father's as well as the mother's education for almost all children, and are able to track family structure from birth through childhood.

We group each parent's education into three categories: less than 12 years of education ($<HS$) corresponding to high school dropouts in the U.S.; high school graduate (HS) which may include some vocational training or two year college; and bachelor's degree graduate or more (BA) corresponding to a degree from a four year college in the U.S. The latter category covers professional bachelor degrees (e.g. school teacher, nursing, physiotherapist, social worker) as well as university and business school degrees.

Our primary measure of family structure is parental marital status at birth. For models using our sample of full siblings, we use parents' marital status at the birth of the youngest of their joint children. We choose this alternative definition of marital status because it is very common in Denmark to marry after the birth of the first child and eventual marital status seems to provide a better indicator of the parental relationship as shared by siblings. As almost all parents with more than one child are either married or cohabiting at the time of the youngest of their joint children, we only distinguish between having married and non-married parents.[7]

For the models of adult outcomes, we consider family structure measured at age 12 as well as parental marital status at birth.[8] For childhood family structure, we distinguish between three types: traditional families where children live with both biological parents (*Trad*), with no distinction between married and cohabiting parents; step-families in which children live with one biological parent and a step-parent (*Step*); and single parent families (*Single*). Using childhood family structure, though it may

[7]Less than two percent of the sibling sample have parents who never cohabit and who are never married at any of the childbirths.

[8]More precisely, family structure at age 12 is measured on January 1^{st} of the year the child turns 13. We also considered family structure at age 16 with very similar results.

be endogenous with respect to child outcomes, allows us to include birth cohorts going back to 1966, while marital status at birth is observed only in the medical birth registry which begins in 1973. Family structure at birth and at age 12 are strongly correlated,[9] and results using both measures are quantitatively similar.

3.2 Outcome Variables: From Adolescence through Adulthood

The outcomes of interest fall into two groups: 1) School outcomes measured in adolescence and 2) Educational attainment and labor market outcomes measured at age 31. Since these outcomes span from age 16 through age 31 (and in some specifications through age 41) and come from several administrative registers, different birth cohorts will be used in analyses of outcomes in adolescence and adulthood; Appendix Table A1 summarizes the cohorts used for each part of the analysis.[10] We have one outcome that is available for all cohorts, completion of grade 9 on time, and we use this outcome to examine whether the gender gap in the effects of family environment has changed over time.

In Denmark, the first nine years of schooling constitute primary school and are mandatory. Children are required to start first grade the year they turn 7, though parents are able to apply for an exemption such that their child starts school a year earlier or later. Boys are about twice as likely to delay school start compared to girls (Dee and Sievertsen, 2015). Grade repetition is very rare; Simonsen et al. (2015) show that on average less than 0.5 percent are retained or delayed for each grade level from grade 1 to 9. Whether the child completes grade 9 on time is a marker of academic achievement that reflects a combination of early school readiness and success in school progression, and is strongly correlated with final educational attainment.

[9] Of those last born children who were born to married (non-married) parents, 80.87 (60.71) percent live in a traditional family at age 12, while 7.48 (14.48) percent live in a step family and 11.65 (24.82) percent live with a single parent.

[10] When we refer to outcomes at a certain age, we always refer to the age the individual turns during the particular year. Thus, grade 9 outcomes are measured at age 15 for about half of the sample, since the school year ends in June.

At the end of primary school, students take the final grade 9 exam, which is the same across the country and is required for all students who continue to academic high school.[11] Our second school outcome is the overall GPA obtained at the end of grade 9 (based on all grades received both from teacher assessment and final exams).[12] Other early outcomes include indicators of having received a diagnosis for behavioral and emotional disorders at a hospital[13] and attending special education during grade 9.[14] Since the administrative data on grade 9 GPA begins in 2002, we consider birth cohorts born from 1986 to 1995 for this part of the analysis.

After primary school, students can choose to continue to academic high school, which takes three years, or vocational training programs of differing lengths (predominantly 4 or 4.5 years). A diploma from the academic high school is necessary to apply for university. A bachelor's degree from university takes three years (i.e. 15 years of completed education) and a master's degree takes two additional years. Instead of university, it is possible for academic high school graduates to take a two year college degree or to enter vocational training.

Figure 1 graphs the share of people who have completed certain levels of education at each age from 15 to 35 years. This shows that by age 31 almost everyone has completed their education, and we study educational attainment and labor market outcomes at this age. Educational attainment is measured as highest completed education measured in months or, alternatively, having at least a high school or BA degree.[15] In terms of labor market outcomes, we consider whether the person is em-

[11]Since 2007, the exam has been mandatory for all students.

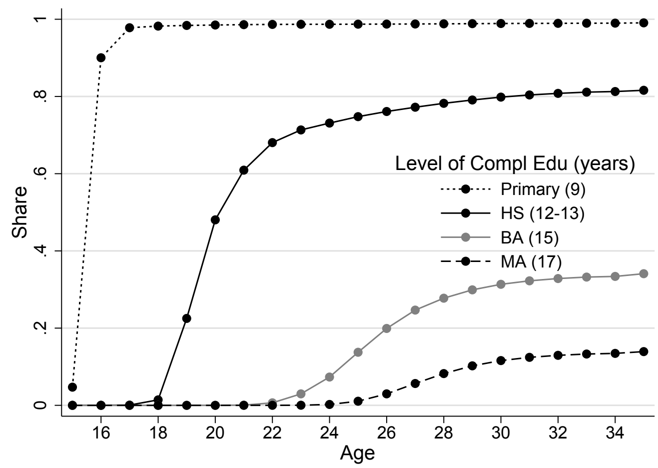
[12]Similar results are found if we instead only use the GPA from written national exams in Math and Danish. If we consider the test scores from these exams separately, the differential effects are mainly found for Danish.

[13]This outcome is defined from hospital records in the Danish national patient registry and is based on both inpatient and outpatient hospital records. We require that the person has received at least one diagnosis by age 21; 92 percent of those with a diagnosis by age 21 have received it by age 16. This measure will be incomplete, since psychiatrists working outside the hospital system do not report to the registry; see the discussion in Obel et al. (2015).

[14]Special education is only observed for 2007-2011.

[15]Highest completed education is measured in months, as some programs take a non-integer number of years. The category *high school* covers academic high school and vocational training with a length of at least 12 years.

Figure 1
Educational Attainment in Denmark by Age



Note: Share of individuals (birth cohorts 1973–1984) with the specified educational level or more at each age from 15–35 years. The category *HS* covers academic high school and vocational training with a length of at least 12 years.

ployed and the annual labor earnings percentile by year of birth and gender.[16] We standardize earnings in this way, following (Chetty et al., 2014), to get a measure of relative income that includes individuals with zero earnings and is comparable across cohorts and gender. For the primary analysis of adult outcomes, the sample consists of individuals born between 1973 and 1984, for whom we can observe both parents' marital status at birth and outcomes at age 31.[17]

4 Sample Selection and Empirical Framework

4.1 Summary Statistics

So that we observe the family environment during childhood as well as adult outcomes, we consider individuals born between 1973 and 1995 for the main analysis; for the

[16]Individuals are defined as being employed if they have any positive labor earnings or have employment as the main source of income, including self-employment. The measure of earnings is the total sum of income earned from wage employment during a particular year.

[17]When we examine whether the gender gaps in family effects have changed over time, we expand the sample to cohorts born between 1966 and 1984.

Table 1
Sample Selection: Averages of Family Environment Variables (Percent)

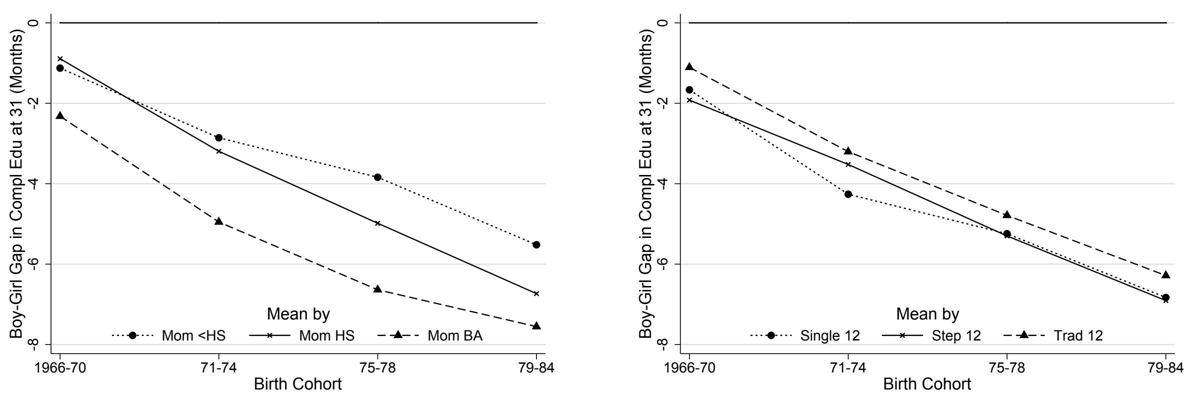
Samples	*a*) Entire Population	*b*) Total Pop for Estimation	*c*) Sibling Sample
Selection Criteria and Background Information			
Child is male	51.21	51.26	51.38
Child's year of birth	1984	1984	1984
Child is twin	2.23	0.00	0.00
Child is adopted	0.93	0.00	0.00
# of Children in Family	2.61	2.55	2.72
Child's birth order	1.81	1.78	1.84
Mother immigrant	9.36	5.31	5.98
Father immigrant	8.97	5.78	6.35
Mother's age at birth (years)	27.05	27.11	26.97
Parental Education			
Mother <HS	34.76	34.15	32.63
Mother HS	37.83	38.07	37.87
Mother BA	27.41	27.78	29.50
Father <HS	26.81	26.52	24.99
Father HS	52.47	52.79	53.20
Father BA	20.71	20.69	21.81
Marital Status			
Married at own birth	63.04	63.53	68.11
Married at youngest sib birth	70.62	71.48	79.47
Family Structure at Youngest Sib Age 12 (Cohorts 1973-84)			
Traditional family	70.80	75.29	80.75
Step-parent family	9.53	9.78	6.96
Single-parent family	19.67	14.93	12.29
N	1,472,204	1,289,542	888,635

The columns represent three different samples of individuals born between 1973 and 1995 who all have an observation on Grade 9 completion on time: a) the entire population; b) the total population (i.e. individuals who are not twins or adoptees and have an observation on all parental variables); and c) the sibling sample (i.e. those who are observed in (b) and have at least one biological sibling in the sample). Note that the following variables are *not* reported as percent: year of birth, # of children in family, and birth order.

analysis of educational attainment across cohorts, we include cohorts going back to 1966. We restrict the sample to those for whom we observe all parental variables[18] and include only families without adopted children and only singleton births.[19] For the main analysis, we consider this sample (referred to as the *total population*) as well as the subsample of families with at least two full siblings (i.e. children with the same mother and same father; referred to as the *sibling sample*).

Figure 2

Gender Gap in Highest Completed Education (in Months) at Age 31



(a) By Maternal Education **(b)** By Family Structure

Note: Male-female gap in highest completed education at age 31 by maternal education and family structure at age 12, respectively. The gap is calculated as the difference between the raw mean for each gender by birth cohort for the total population.

Table 1 provides descriptive statistics for three different samples of individuals born between 1973 and 1995 who all have an observation on Grade 9 completion on time: a) the entire population[20]; b) the total population sample; and c) the sibling sample. The largest differences between the sibling sample and the other two samples are that, on

[18]Since parental education is a key variable for the analysis, we restrict the sample to those families where we observe both parents' education. Mother's (father's) education is missing for 1.28 (2.96) percent of children who would otherwise have been in the total population sample. Results from models including only mother's education are not sensitive to excluding or including the children without information on father's education.

[19]However, as a robustness check, we have also estimated the core models in a sample of gender-discordant twins and get consistent results.

[20]Not reported, the percent of missing observations for the following variables are: mother's age 0.32, marital status 4.75, father's identity 2.12, mother's education 2.15, and father's education 5.01.

average, children in the former sample have slightly better educated parents, parents more likely to be married at birth, and slightly larger families. In the total population and the sibling sample, children are less likely to have immigrant parents than in the entire population because some children of immigrant parents are immigrants themselves and therefore not observed at birth.[21]

Figure 2 shows the raw gender gap in educational attainment at age 31 by birth cohort and childhood family environment.[22] Educational attainment was almost equal for men and women born in the first period (1966–1970). For subsequent cohorts, the gender gap has increased such that women born between 1979 and 1984 have attained on average six months more education by age 31 than their male counterparts. The educational gender gap is smallest for the children of less-educated mothers. In contrast, there is little variation in the gender gap by family structure.

4.2 Empirical Framework

Our goal is to identify whether childhood family environment has a differential impact on men's adult outcomes relative to women's to test the "vulnerable boys" hypothesis. The empirical strategy is twofold: First, we compare the male-female differences in adolescent and adult outcomes by family environment (i.e. a difference-in-difference strategy). Second, we focus on differences between brothers and sisters with the same mother and father by family characteristics (i.e. controlling for family fixed effects).

We begin by estimating an OLS model of the effect of family environment on outcomes for boys and girls using the total population sample. The outcome Y of individual i in family j exposed to the family environment, *Fam Env*, is given by:

$$Y_{ij} = \beta_0 + \beta_1 Male_i + \beta_2 Fam\ Env_i + \beta_3 Male_i \times Fam\ Env_i + X_i'\theta + \nu_{ij}, \qquad (1)$$

[21]The primary reason for a missing observation on marital status at birth is because the child is not observed in the birth registry (i.e. is born outside Denmark). Appendix Table A2 shows descriptives statistics by cohort for the total population and the sibling sample.

[22]The raw gender gap by paternal education is very similar to the one observed in Figure 2; the gender gap for the sibling sample is similar as well.

where X_i is a vector of individual controls (year and month of birth, birth order, mother's age at birth, family size, and parental immigrant status)[23] and standard errors, ν_{ij}, are clustered at the family level. For the estimation, we exploit the randomness in child gender; as long as child gender is independent of family environment, β_3 represents the causal effect of family environment on gender differences in adult outcomes.[24]

However, these estimates may be biased if family structure and child gender are not independent. Sex-selective abortion, which might generate a correlation between marital status and child gender, is not expected to be an important consideration in the Danish context, but there is considerable evidence from a number of countries that fathers are more likely to co-reside with, seek custody of, and marry the mothers of their sons rather than daughters (Lundberg and Rose, 2003; Dahl and Moretti, 2008; Lundberg, 2005). There is also increasing evidence that the Trivers-Willard hypothesis, which suggests that females in advantaged circumstances are more likely to bear male offspring, may apply to human populations through the impact of stress on the mortality of male and female fetuses (Almond and Edlund, 2007; Hamoudi and Nobles, 2014; Norberg, 2004; Trivers and Willard, 1973); though the effects of even extreme events are small. If these factors generate systematic selection of boys and girls across family types, cross-sectional models of the effects of family environment will be misleading.

To consider whether selection into specific family types by gender might be a problem, Table 2 considers whether observed family characteristics differ for boys and girls. This is done by regressing an indicator for being male on family characteristics on the total population sample. Men have on average more siblings than women and are more likely to live with both biological parents at age 12. The coefficients on parental education and marital status at birth are not jointly, significantly different from zero

[23]For the adult outcomes, we also include dummies for age at observation if not observed at age 31 (age 29 or 30).

[24]To ease interpretation, we multiply all coefficients on binary variables (as well as the grade 9 GPA and earnings percentile) by 100 such that the interpretation of, for instance, β_3 becomes a change of β_3 percentage points instead of $\beta_3 \times 100$ percentage points.

Table 2
Balancing of Characteristics by Gender

	Dependent Variable: Being Male					
Birth Cohorts	1973–1995				1973–1984	
	(1)		(2)		(3)	
# of Children in Family	0.25***	(0.04)	0.25***	(0.04)	0.24***	(0.06)
Mother's age at birth	-0.01	(0.01)	-0.01	(0.01)	-0.00	(0.01)
Mother immigrant	-0.23	(0.24)	-0.26	(0.24)	0.08	(0.36)
Father immigrant	-0.27	(0.23)	-0.28	(0.23)	-0.29	(0.35)
Mother HS	-0.02	(0.11)	-0.03	(0.11)	-0.02	(0.15)
Mother BA	0.11	(0.13)	0.11	(0.13)	0.35**	(0.18)
Father HS	0.08	(0.11)	0.07	(0.11)	0.22	(0.15)
Father BA	0.25*	(0.14)	0.24*	(0.14)	0.19	(0.20)
Married at own birth	-0.13	(0.10)				
Married at y. sib birth			0.01	(0.10)		
Trad family age 12					0.61***	(0.17)
Step family age 12					0.45*	(0.25)
N	1,289,542		1,289,542		650,633	
Mean of Y	51.26		51.26		51.11	
$Prob > F1$	0.00		0.00		0.00	
$Prob > F2$	0.14		0.24		0.00	

Standard errors in parentheses. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The sample is the total population. The dependent variable is an indicator for being male. All regressions control for year and month of birth dummies and a constant. $Prob > F1$ is the p-value for a joint F-test of whether all the shown estimates are jointly equal to zero. $Prob > F2$ is the p-value for a joint F-test of whether all estimates of parental education and family structure are jointly equal to zero.

($Prob > F2$), though men are slightly more likely to have BA educated parents.

As an alternative empirical approach, we focus on differences between brothers and sisters with the same mother and father who were raised in the same household:

$$Y_{ij} = \alpha_0 + \alpha_1 Male_i + \alpha_2 Male_i \times Fam\ Env_j + X_i'\gamma + \mu_j + \epsilon_{ij}, \tag{2}$$

where μ_j is a family fixed effect. In this model, α_2 represents the causal effect of family environment on gender differences in adult outcomes as long as we do not omit any important time-varying variable.[25] This empirical strategy has drawbacks, however. In addition to cutting the sample size in half, it restricts the source of variation to families with at least two gender-discordant siblings. If there are behavioral spillovers between siblings, or if patterns of parental investments are different in only-child families or families with same-sex children, then the estimate of α_2 from the sibling sample may not be representative of the effects of family environment in all families.[26]

To examine whether the gender gaps in the effects of family environment have changed across birth cohorts, we also interact the independent variables of interest with a vector of birth cohort dummies (grouped into intervals), C:

$$Y_{ij} = \delta_0 + (Male_i \times C_i)'\delta_1 + (Male_i \times Fam\ Env_j \times C_i)'\delta_2 + (Fam\ Env_j \times C_i)'\delta_3 + X_i'\eta + \mu_j + \zeta_{ij}. \tag{3}$$

[25] As a robustness check (Appendix Table A8), we restrict the sample to only those full siblings who experience the same observable childhood family structure at birth.

[26] Interpreting the coefficients β_3 and α_2 in equations (1) and (2) as indicators of the causal effect of post-natal family environment on child development requires that we assume that parental resources do not affect the endowments of boys and girls at birth differently. Autor et al. (2016) show that this assumption is reasonable for their Florida data; they do not find a sibling gender gap in the effects of family conditions on birth outcomes when controlling for mother fixed effects. In Appendix Table A3, we show that the prenatal inputs and birth outcomes of sisters and brothers are also not differently affected by the family environment in the Danish data.

5 Results

5.1 Outcomes in Adolescence

Table 3 reports key coefficients from the models of Grade 9 outcomes, using both the total population and sibling samples. Column (1) shows that, for the total population, boys are 9.3 percentage points less likely to complete grade 9 on time than girls, conditional on year and month of birth, birth order, maternal age at birth, family size, and parental immigrant status.[27] Column (2) adds mother's education and marital status at birth as well as interaction terms between these variables and a male dummy. Boys benefit more from having a highly educated mother (HS and BA degree) compared to girls; the male disadvantage is reduced by 0.8 and 1.8 percentage points for boys of HS and BA educated mothers, respectively. Males also benefit from being born to married parents. Column (3) adds father's education and highlights one advantage of our data (i.e. that we observe fathers' characteristics for almost all children): the benefit of mother's education for boys diminishes substantially when father's education is included and father's education further reduces the gender gap. For highly educated fathers, the gender gap is reduced by 2.5 percentage points (22 percent) compared to children with less than HS educated fathers. Column (4) estimates the same model as in Column (3) but on the sibling sample rather than the total population with very similar point estimates and significance levels. Finally, Column (5) includes family fixed effects for the sibling sample, which again give very similar results compared to using the total population without fixed effects.

The gender gap in grade 9 GPA is large —almost 30 percent of a standard deviation [Column (6)]. The results for this outcome are somewhat different from other adolescent outcomes in that some indicators of parental resources increase, rather than decrease, the gender gap and there are some discrepancies between the results from the total and the sibling samples. Paternal college education reduces the gender gap

[27] This number is 8.9 percentage points for the sibling sample without controlling for family fixed effects and 9.1 percentage points with fixed effects, see Columns (3) and (6) in Panel A in Appendix Table A9.

in grade 9 GPA [Columns (8) to (10)]; this is true for all model versions and the effect nearly doubles in the fixed-effect model. However, having married parents at birth increases the gap. Maternal education and father's HS education also increase the gender gap in the OLS estimates, but are insignificant in the fixed effects model. To the extent that parents, particularly mothers, influence their children's grades, there may be equalizing forces within households with gender-discordant siblings that account for these different results.

Table 3

Grade 9 On Time and GPA

	Grade 9 On Time					Grade 9 GPA				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Male	-9.30***	-10.90***	-11.42***	-11.28***	-11.03***	-29.64***	-27.73***	-27.07***	-29.31***	-31.11***
	(0.09)	(0.22)	(0.26)	(0.38)	(0.48)	(0.26)	(0.56)	(0.67)	(1.01)	(1.10)
Male×Mom HS		0.82***	0.56**	0.89***	0.80*		-3.19***	-3.00***	-1.87**	-1.20
		(0.24)	(0.24)	(0.32)	(0.41)		(0.64)	(0.64)	(0.90)	(0.96)
Male×Mom BA		1.78***	0.83***	0.80**	1.07**		-0.81	-2.26***	-0.43	1.28
		(0.25)	(0.27)	(0.35)	(0.45)		(0.67)	(0.72)	(0.98)	(1.06)
Male×Dad HS			0.80***	0.79**	0.65			-1.98***	-1.66*	0.39
			(0.24)	(0.32)	(0.41)			(0.64)	(0.90)	(0.96)
Male×Dad BA			2.53***	2.45***	1.71***			2.32***	2.53**	4.95***
			(0.29)	(0.39)	(0.49)			(0.80)	(1.08)	(1.16)
Male×Married		1.34***	1.22***	1.13***	0.76**		-1.62***	-1.65***	-1.40*	-1.35*
		(0.18)	(0.18)	(0.26)	(0.34)		(0.51)	(0.50)	(0.74)	(0.80)
N	579,049	579,049	579,049	335,241	335,241	524,330	524,330	524,330	288,667	288,667
Mean of Y	82.45	82.45	82.45	83.38	83.38	-1.41	-1.41	-1.41	2.98	2.98
Total Population OLS	X	X	X			X	X	X		
Sibling Sample OLS				X					X	
Sibling Sample FE					X					X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The samples consist of individuals born from 1986–1995. *Grade 9 on time* indicates whether the person completed grade 9 by age 16. *GPA* is an average of all grades given during grade 9 both from teacher assessment and final exams for all subjects and is standardized with mean zero and standard deviation of one by year of grade 9 completion for the entire population. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), and a constant. All OLS models also control for family size dummies, parental immigrant status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table 4
Behavioral & Emotional Disorder and Special Education

	Behavioral & Emotional Disorder					Special Education				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Male	0.96***	1.22***	1.40***	1.22***	1.18***	1.36***	2.37***	3.03***	3.55***	3.10***
	(0.03)	(0.08)	(0.10)	(0.14)	(0.18)	(0.05)	(0.16)	(0.19)	(0.34)	(0.45)
Male×Mom HS		-0.03	0.03	0.09	-0.09		-0.82***	-0.64***	-1.00***	-1.06***
		(0.09)	(0.09)	(0.11)	(0.15)		(0.16)	(0.16)	(0.27)	(0.36)
Male×Mom BA		-0.33***	-0.18*	-0.13	-0.20		-1.44***	-1.06***	-1.17***	-1.13***
		(0.09)	(0.10)	(0.12)	(0.16)		(0.16)	(0.16)	(0.27)	(0.37)
Male×Dad HS			-0.30***	-0.29**	-0.18			-1.07***	-1.58***	-1.28***
			(0.09)	(0.12)	(0.15)			(0.16)	(0.27)	(0.35)
Male×Dad BA			-0.44***	-0.44***	-0.36**			-1.32***	-1.89***	-1.58***
			(0.11)	(0.13)	(0.18)			(0.17)	(0.28)	(0.37)
Male×Married		-0.26***	-0.24***	-0.09	-0.07		-0.30***	-0.24**	-0.10	0.23
		(0.07)	(0.07)	(0.09)	(0.13)		(0.10)	(0.10)	(0.18)	(0.25)
N	579,049	579,049	579,049	335,241	335,241	284,769	284,769	284,769	95,126	95,126
Mean of Y	1.63	1.63	1.63	1.45	1.45	1.92	1.92	1.92	1.63	1.63
Total Population OLS	X	X	X			X	X	X		
Sibling Sample OLS				X					X	
Sibling Sample FE					X					X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The samples consist of individuals born from 1986–1995. *Behavioral & Emotional Disorder* indicates whether the person has been diagnosed with ICD-10 codes F90-98 by the age of 21 years at a hospital. *Special education* indicates whether the person attends special education during grade 9 and is only observed for years 2007-2011. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), and a constant. All OLS models also control for family size dummies, parental immigrant status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table 4 examines sibling differences in outcomes that reflect a combination of behavioral and developmental problems —diagnosis for behavioral and emotional disorder and attending special education during 9^{th} grade. From Columns (1) and (6), it is clear that these outcomes are more prevalent among boys: 66 percent of those with behavioral and emotional problems and 69 percent of those attending special education are male.[28] The OLS models of behavioral and emotional disorder indicate that higher parental education reduces the gender gap [Columns (2) to (4)] in the total sample. Only father's BA is still significant in the sibling model with fixed effects. However, both maternal and paternal education decrease the probability of attending special education much more for boys than girls [Columns (5) to (6)], and this effect is consistent across samples and models.

We observe on-time completion of grade 9 consistently across cohorts, and therefore use this outcome to look at whether the gender gap in the effects of family environment has changed over time. Since we examine adult outcomes as well, we want to know whether any variation in the effects of family background between childhood and adult outcomes are due to different ages at observation or due to different birth cohorts.

Table 5 shows the male-cohort-family environment interactions from estimates of equation (3) with the outcome grade 9 completion on time. The base male-female gap in grade 9 on time has increased modestly over time [Column (1)]; boys in the omitted group (i.e. with low educated, unmarried parents) born from 1973 to 1978 were 8.4 percentage points less likely to complete grade 9 on time compared to girls, and this gap increased to more than 11 percentage points for those born between 1990 and 1995. In contrast, the male premium in the effects of maternal and paternal education has been relatively stable over this time period [Columns (2) to (5)]; the estimated effects do not differ significantly across cohorts. This indicates that the gender gap in the effects of parental education is not a recent phenomenon, but has been relatively constant over more than two decades. This suggests in turn that the effects on adult

[28] Kristoffersen et al. (2015) find a strong association between behavioral problems and school outcomes for Danish children, but the behavioral gender gap explains only a fraction of the gender difference in test scores.

outcomes that we will observe for older cohorts may be predictive of the experiences of more recent cohorts. Meanwhile, the male advantage of being born to married parents has declined and is not significant for the youngest cohort [Column (8)], which may indicate that the role of cohabitation has changed over time as well.

Table 5

Grade 9 On Time by Birth Cohort: Male-Cohort-Family Environment Interactions

	Male (1)	Mom HS (2)	Mom BA (3)	Dad HS (4)	Dad BA (5)	Married (6)
Male×1973-78	-8.42***	1.54***	2.08***	0.84**	1.64***	1.93***
	(0.68)	(0.38)	(0.45)	(0.40)	(0.50)	(0.63)
Male×1979-84	-9.18***	0.71*	1.28***	0.67	2.29***	2.04***
	(0.54)	(0.39)	(0.43)	(0.41)	(0.51)	(0.46)
Male×1985-89	-10.92***	1.11**	1.69***	0.81*	2.20***	1.07**
	(0.56)	(0.46)	(0.50)	(0.48)	(0.57)	(0.43)
Male×1990-95	-11.22***	1.00**	1.33**	0.67	1.63***	0.66
	(0.57)	(0.49)	(0.53)	(0.49)	(0.59)	(0.41)
$Prob > F1$	0.00	0.50	0.58	0.99	0.72	0.10
$Prob > F2$	0.00	0.39	0.29	0.79	1.00	0.09

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The sample used for estimation is the sibling sample with individuals born from 1973–1995. $N = 888,635$. The outcome is Grade 9 on time with a mean of 86.68 percent. All estimates come from one regression as specified in equation (3), i.e. a regression interacting the male-family environment interactions as well as the family environment variables with birth cohort dummies. The model also controls for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), and a constant. $Prob > F1$ reports the p-value from a joint F-test of whether all the estimates in the particular column are equal. $Prob > F2$ reports the p-value from a joint F-test of whether the estimates for the earliest and the last cohorts in the particular column are equal.

In general, our results on school and behavioral outcomes in adolescence are consistent with previous studies finding that boys benefit more from a good family background than girls in terms of outcomes indicative of learning and developmental problems (Bedard and Witman, 2015; Bertrand and Pan, 2013; Autor et al., 2016). Using Danish data, we are able to support the overall finding for the U.S. that boys seem more vulnerable to a disadvantageous family environment than their sisters when looking at adolescent outcomes. Notably, we find that boys benefit differentially from high paternal education, and that the effects of parental education have been relatively stable over time.

5.2 Adult Outcomes

When we turn to educational attainment, employment, and earnings at age 31 we find, in contrast to school-age outcomes, that women benefit more from higher maternal education than men. This is true for both samples and is robust to the inclusion of family fixed effects. For the total population, men complete less education than women with a raw gender gap of 5.4 months [Table 6, Column (1)]. This gap is strongly increasing in maternal education [Column (2)]. The gender gap in educational attainment rises from 4.9 months for children of less than HS mothers to 6.0 and 7.7 months for children with, respectively, HS and BA educated mothers [Column (2)], holding marital status constant. Column (3) adds father's education and its interaction with the male dummy instead of mother's education. These results suggest that women benefit more from father's BA education than men. However, when including both parents' education in Column (4), it becomes clear that men benefit more than women from paternal education while the opposite is true for maternal education. These results are insensitive to the inclusion of family fixed effects [Column (6)]. For the sibling sample controlling for family fixed effects, the results show that having a HS or BA educated father reduces the gender gap by 0.8 and 1.5 months, respectively, while having a HS or BA educated mother increases the gender gap by 1.1 and 2.8 months.[29] For the total population, having married parents at birth decreases the gender gap by 0.7 months. In contrast, neither marital status at birth nor childhood family structure significantly affect the gender gap in educational attainment for the sibling sample. Overall, these results indicate that women's educational attainment is more responsive to maternal education, while men benefit more from paternal education.

[29] In results not reported here, we have also considered the natural logarithm of educational attainment to examine whether we would find a similar pattern when looking at relative instead of absolute differences. Those results are in line with the ones reported here on educational attainment.

Table 6

Educational Attainment at Age 31: Highest Completed Education (Months)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Male	-5.40***	-4.86***	-5.79***	-5.09***	-5.16***	-4.98***	-5.05***
	(0.07)	(0.15)	(0.16)	(0.17)	(0.30)	(0.36)	(0.36)
Male×Mom HS		-1.17***		-1.30***	-1.36***	-1.11***	-1.13***
		(0.15)		(0.15)	(0.21)	(0.25)	(0.25)
Male×Mom BA		-2.62***		-3.00***	-2.98***	-2.82***	-2.83***
		(0.17)		(0.19)	(0.25)	(0.30)	(0.30)
Male×Dad HS			-0.05	0.33**	0.70***	0.79***	0.77***
			(0.16)	(0.16)	(0.21)	(0.25)	(0.25)
Male×Dad BA			-0.43**	1.04***	1.23***	1.52***	1.50***
			(0.20)	(0.22)	(0.29)	(0.34)	(0.34)
Male×Married		0.69***	0.66***	0.66***	0.43	0.20	
		(0.15)	(0.15)	(0.15)	(0.28)	(0.34)	
Male×Trad 12							0.37
							(0.34)
Male×Step 12							-0.23
							(0.49)
N	632,508	632,508	632,508	632,508	355,090	355,090	355,090
Mean of Y	162.97	162.97	162.97	162.97	164.14	164.14	164.14
Total Pop. OLS	X	X	X	X			
Sibling Sample OLS					X		
Sibling Sample FE						X	X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The samples consist of individuals born from 1973–1984. *Highest completed education* measures the length of highest completed education in months. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant. All OLS models also control for family size dummies, parental immigration status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table 7

Level of Educational Attainment at Age 31: Having at least a HS and BA Degree

	HS Degree or more					BA Degree or more				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Male	-1.97***	-1.96***	-1.99***	-1.81***	-2.34***	-16.36***	-12.75***	-12.37***	-12.03***	-10.92***
	(0.09)	(0.28)	(0.49)	(0.60)	(0.59)	(0.12)	(0.25)	(0.45)	(0.57)	(0.55)
Male×Mom HS		-1.30***	-1.41***	-1.41***	-1.47***		-3.41***	-3.56***	-2.96***	-2.87***
		(0.23)	(0.30)	(0.37)	(0.37)		(0.26)	(0.35)	(0.43)	(0.43)
Male×Mom BA		-2.07***	-1.98***	-2.02***	-2.05***		-7.17***	-7.26***	-7.03***	-6.97***
		(0.25)	(0.32)	(0.40)	(0.40)		(0.33)	(0.44)	(0.54)	(0.54)
Male×Dad HS		0.88***	1.25***	1.63***	1.58***		-2.15***	-1.92***	-2.09***	-2.00***
		(0.24)	(0.32)	(0.40)	(0.40)		(0.25)	(0.34)	(0.42)	(0.42)
Male×Dad BA		0.61**	0.58	1.02**	0.95**		3.04***	3.32***	3.41***	3.54***
		(0.28)	(0.36)	(0.45)	(0.45)		(0.37)	(0.50)	(0.61)	(0.61)
Male×Married		0.47**	0.55	0.20			-0.26	-1.28***	-1.47***	
		(0.22)	(0.42)	(0.53)			(0.24)	(0.45)	(0.56)	
Male×Trad 12					1.01*					-3.17***
					(0.52)					(0.53)
Male×Step 12					-0.33					-0.42
					(0.80)					(0.76)
N	632,508	632,508	355,090	355,090	355,090	632,508	632,508	355,090	355,090	355,090
Mean of Y	82.17	82.17	83.49	83.49	83.49	35.04	35.04	36.58	36.58	36.58
Total Population OLS	X	X				X	X			
Sibling Sample OLS			X					X		
Sibling Sample FE				X	X				X	X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The samples consist of individuals born from 1973–1984. *HS Degree* is an indicator for whether the individual has completed at least 12 years of education. *BA graduate* indicates whether the person has at least a BA degree. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant. All OLS models also control for family size dummies, parental immigration status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table 8

Labor Market Outcomes at Age 31

	Employed					Earnings Percentile by Birth Cohort and Gender				
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
Male	2.86***	4.79***	4.76***	5.49***	5.64***	-0.44***	0.67***	0.86***	1.23***	1.00***
	(0.08)	(0.22)	(0.39)	(0.51)	(0.50)	(0.07)	(0.17)	(0.30)	(0.39)	(0.38)
Male×Mom HS		-1.48***	-1.70***	-1.80***	-1.76***		-0.97***	-0.90***	-1.22***	-1.23***
		(0.18)	(0.23)	(0.30)	(0.30)		(0.16)	(0.21)	(0.27)	(0.27)
Male×Mom BA		-1.75***	-1.80***	-1.96***	-1.95***		-1.99***	-2.15***	-2.44***	-2.46***
		(0.22)	(0.28)	(0.36)	(0.36)		(0.20)	(0.26)	(0.33)	(0.33)
Male×Dad HS		-1.39***	-1.50***	-2.19***	-2.16***		-0.74***	-0.80***	-1.19***	-1.22***
		(0.19)	(0.24)	(0.32)	(0.32)		(0.16)	(0.21)	(0.27)	(0.27)
Male×Dad BA		-2.08***	-2.28***	-2.30***	-2.27***		-2.50***	-2.54***	-2.58***	-2.62***
		(0.25)	(0.32)	(0.42)	(0.42)		(0.23)	(0.30)	(0.38)	(0.38)
Male×Married		0.21	0.40	0.25			0.81***	1.02***	1.25***	
		(0.18)	(0.35)	(0.46)			(0.15)	(0.29)	(0.36)	
Male×Trad 12					-0.04					1.67***
					(0.44)					(0.35)
Male×Step 12					0.80					0.65
					(0.67)					(0.52)
N	643,219	643,219	365,676	365,676	365,676	643,219	643,219	365,676	365,676	365,676
Mean of Y	89.54	89.54	90.18	90.18	90.18	53.99	53.99	54.64	54.64	54.64
Total Population OLS	X	X				X	X			
Sibling Sample OLS			X					X		
Sibling Sample FE				X	X				X	X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The samples consist of individuals born from 1973–1984. *Employed* takes the value one if the person has positive labor earnings or have employment as the main source of income including self-employment and zero otherwise. *Earnings percentile* measures the annual earnings percentile by gender and year of birth. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant. All OLS models also control for family size dummies, parental immigration status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Turning to the binary outcomes of having received a HS or BA degree by age 31, we find results quite similar to those in the educational attainment model, with two exceptions for the probability of receiving a BA degree. First, women of HS educated fathers benefit more than their brothers, though boys benefit more from a college-educated father. Second, women in the sibling sample also benefit more than their brothers from having parents who were married at the youngest sibling's birth [Columns (8) to (9)] and living in a traditional family during childhood [Column (10)]. These results are sharply at odds with those we saw for school outcomes at age 16: on most dimensions, women benefit more from a favorable childhood family environment than their brothers in terms of higher educational attainment in adulthood.

Table 8 presents results for labor market outcomes at age 31. Column (1) shows that men are 2.9 percentage points more likely to be employed than women. However, we again see the pattern that women benefit more than their brothers from having parents with at least HS education, with the effects of paternal education slightly larger than those for maternal education. Parental education also differentially increases the earnings percentile of women relative to their brothers. Women whose parents have at least a HS degree earn more than their brothers relative to their birth cohort and gender, and the effects of maternal and paternal college education are particularly strong (2.3 and 2.7 percentage points). We find no gender gap in the effects of childhood family structure on employment, but men of married parents at birth benefit more than their sisters in terms of earnings.

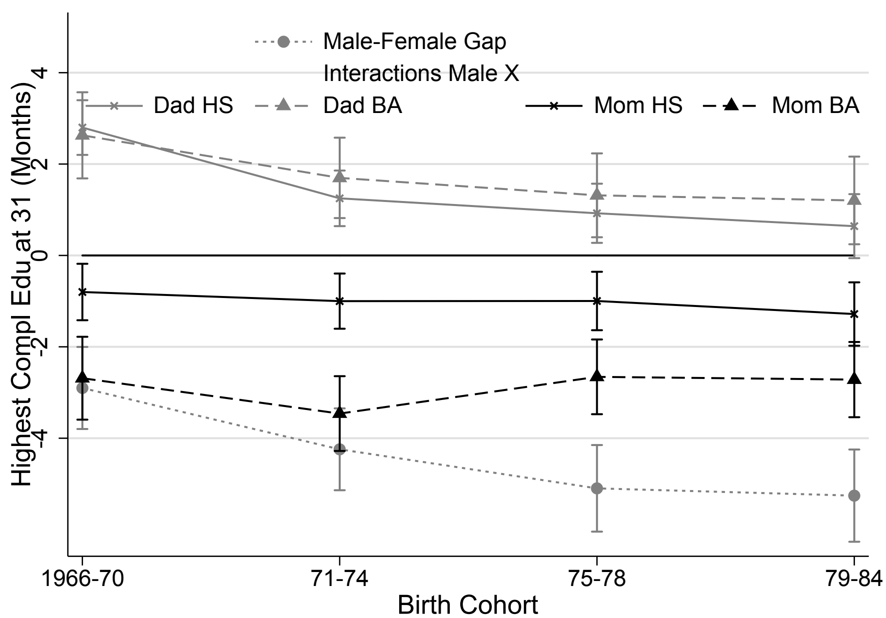
Finally, Appendix Table A4 examines whether the gender gap in the effects of childhood family environment on labor market outcomes vary across ages (26, 31, 36, and 41) by using the sibling sample of individuals born between 1966 and 1971. The results across these four ages show that the differential effects of mother's education on women's employment persist as they age. The age pattern of parental education effects on the earnings percentile is different. At age 26 and 31, both maternal and paternal education have more positive effects on women's earnings. By age 36, parental education no longer has a differential effect on sons and daughters. This change may reflect the different career lifecycles of men and women, especially related to childbirth and household responsibilities (Kleven et al., 2016). Moreover, the male earnings boost from having lived in a traditional family at age 12 grows with age in these older cohorts.

5.3 Educational Attainment Across Cohorts

We have found that women consistently benefit more from high maternal education than their brothers in terms of adult outcomes at age 31. In this subsection, we examine whether these gender differences in the effects of parental education on educational attainment have changed across cohorts.

Figure 3

Highest Completed Education (in Months) at Age 31 by Birth Cohort
Male-Mother's and Male-Father's Education Interaction



Note: The whiskers represent the 95 percent confidence interval. The sample is the sibling sample with individuals born from 1966–1984. All estimates come from one regression interacting the male-family environment interactions as well as the family environment variables with birth cohort dummies. Appendix Table A5 displays all the male-cohort-family environment interaction estimates. The model also controls for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant.

Figure 3 plots the estimates from an educational attainment regression that now includes interactions between the male dummy, family environment, and cohort dummies. From this, we get two important insights. First, it is evident that earlier cohorts of men benefited more from high paternal education than women but that the gender difference in the effect of father's education has diminished substantially and is only borderline significant for recent cohorts. Results reported in the Appendix show a similar pattern of results for receiving a HS and BA degree and also show that the impacts of family structure have been reasonably consistent over time [Appendix Table A5].[30] Family structure has no consistent effect on the gender gap in educational attainment, though living in a traditional family during childhood strongly favors women in all cohorts in terms of the likelihood of receiving a BA. Consequently, this evidence does not support the hypothesis that the increasing prevalence of non-traditional family arrangements explain the growing education gap in favor of girls.

The results in this section show that gender differences in the effects of parental education have been fairly constant across cohorts in terms of educational attainment with one exception: the differentially positive effect of father's education on boys has decreased over time. The same is true of labor market outcomes: the gender specific responses to childhood family environment have been consistent across cohorts, though the more positive effects of both maternal and paternal education on women's employment and earnings tends to diminish with age.

5.4 Sensitivity Analyses

In this subsection, we study the robustness of our findings in three different ways. First, we examine whether different aspects of childhood family environment interact in important ways by gender. Second, we check the robustness of our measure of childhood family structure. Since the main results were quite similar for the different models, we perform these two robustness analyses on the sibling sample including family fixed

[30]The results in Figure 3 are basically identical when only considering traditional families. Though we do not report the results here, if we exclude the male-paternal education interactions, the gender gap in the effects of maternal education appear to be increasing over time, generating a spurious trend in the impact of mother's education.

effects. Third, we compare our main results (estimated on the total population and the sibling sample) to the estimated effects for one-child families and subsamples of the sibling sample divided by the gender composition of the siblings in the sample without family fixed effects.

Appendix Table A6 includes interactions between mother's and father's education in several key models of school and adult outcomes.[31] We find some heterogeneity in the effects of parents' education on educational attainment at age 31.[32] The results suggest that in families where both parents have BA education, men do not benefit more than their sisters from mother's BA education in terms of completed education. For college graduation, the excess female advantage from parental BA education is smaller in families where one parent has BA education and the other has at least HS education. Appendix Table A7 expands the main model by interacting the family environment-male interactions with marital status at birth, but we find little evidence of heterogeneity.

Appendix Table A8 tests the sensitivity of our definition of childhood family structure for the sibling sample, which is based on the experience of the youngest sibling. The results are very robust to using family structure defined for the oldest sibling instead. Alternatively, we restrict the sample to those families with children with the same observed family structure at age 12; the results are again very similar to the main results.

Finally, Appendix Tables A9 and A10 compare OLS models of key outcomes for alternative samples—the total population, children from one-child families, and the samples of full siblings, same-sex siblings and mixed-sex siblings. Overall, the estimates for the different subsamples and for the total population and the sibling sample without fixed effects are similar (both in terms of magnitude and significance), though fewer estimates are significant in smaller samples.

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[31] The correlation between mother's and father's length of education is around 0.41 and 52 percent of parents have the same educational level.

[32] Formally, we test this with an F-test of whether the additional Male$\times$Mom Edu$\times$Dad Edu interaction terms are jointly equal to zero.

6 Conclusion

Motivated by previous findings showing that school-aged boys appear more vulnerable to family disadvantage than school-aged girls, we examine whether such differences persist into adulthood. We use Danish administrative register data, allowing us to examine a broad range of school and adult outcomes for complete cohorts, as well as large samples of full siblings. An advantage compared to previous studies is that we observe both mother's and father's education as well as family structure at birth and during childhood.

In line with findings from the U.S. (Autor et al., 2016; Lundberg, 2016), we first show that in the Danish context boys also appear to be generally more sensitive than girls to family environment in terms of observable outcomes during school. We find the opposite for adult outcomes, including educational attainment, college graduation, employment, and earnings. Women consistently benefit more from maternal education relative to their brothers in terms of education and employment. Paternal education decreases the gender gap in education (favoring sons), though the impact is small. In contrast, paternal education has larger positive effects on the employment and earnings of daughters. Gender gaps in the effects of family structure vary across outcomes, with married parents having significant effects in some samples on college graduation (favoring women) and on high school graduation and earnings percentile (favoring men). Similar results in OLS models using the entire population and family fixed-effect models using a sample of full siblings indicate that selection of boys and girls across family types is not a serious source of bias. Any gender targeting in programs designed to assist students from disadvantaged families should consider the possibility that the effects of deprivation may be more visible for school-age boys than for girls.

Moreover, we show the gender gap in the effects of parental education on completing grade 9 on time has been relatively constant over more than two decades, so the gender difference in the effects of maternal education on primary school completion is not a recent phenomenon. In terms of educational attainment in adulthood, we find that men used to benefit more from paternal education than women but that the gender difference in the effect of father's education disappeared for cohorts born after the mid-1970s. The female premium in the effects of mother's education has been constant for all cohorts.

Although boys respond differently to parental resources and family structure than do girls, the evidence shows that such gender differences do not conform to the simple story that the skill development of boys is particularly sensitive to family environment. Neither can the changes in family structure, in the Danish context, explain the growing education gap in favor of girls. Our findings are compatible with a story in which parental education and other family resources have a strong moderating effect on the observable consequences of behavioral and developmental problems in school that are much more typical of boys than girls. These controlling forces may become less effective as the children become adults, and the results show no indication that long-term skill acquisition, and therefore educational attainment or adult earnings, is affected by these early deficits.

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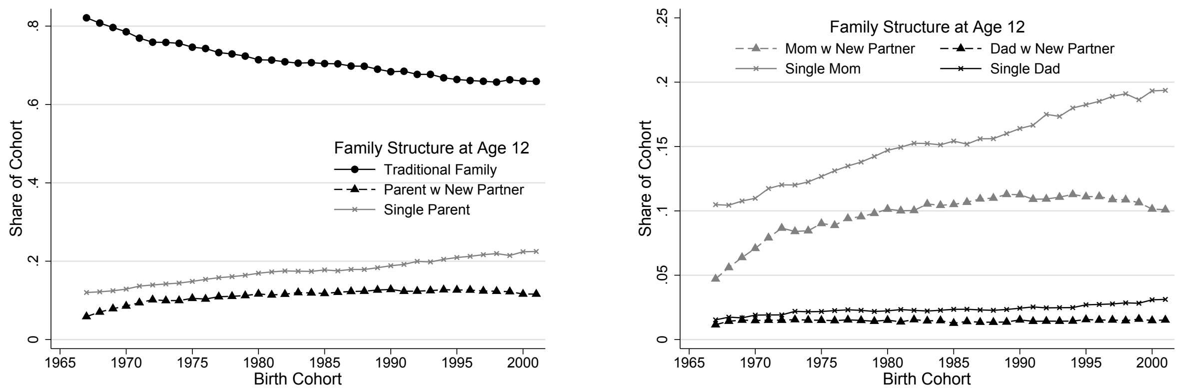
A Appendix

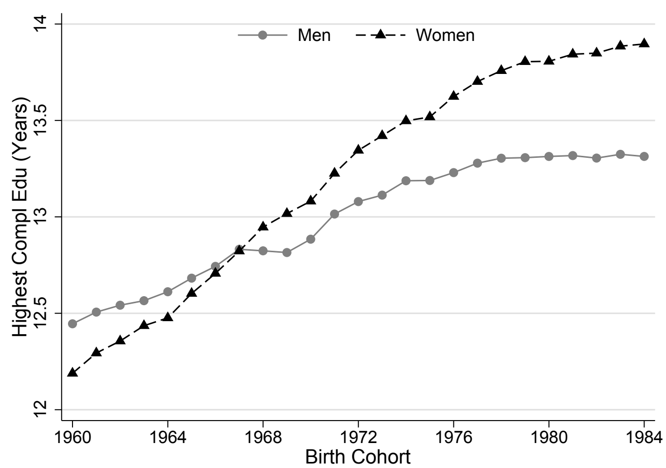
Figure A1
Family Structure at Age 12 (Birth Cohorts 1967–2001)



(a) Traditional, Single, and Step Families **(b)** Non-Traditional Families: Mother vs Father

Note: Family structure as measured on January 1^{st} at age 12 (the year the child turns 13 years). The sample consists of all children living in Denmark at age 12. *Traditional* refers to families in which children live with both biological parents, *Single* refers to families in which children live with only one parent without a new partner, and *Parent w New Partner* refers to families in which children live with one parent and this parent's new partner.

Figure A2
Highest Completed Education (Years) at Age 31 by Gender in Denmark



Note: Highest completed education in years by birth cohort (1960–1984) and gender for people living in Denmark.

Table A1
Overview of Samples and Outcomes

	Main Analysis	Across Cohorts/Ages
Outcomes in Adolescence		
Grade 9 On Time	1986–1995	1973–1995
Grade 9 GPA	1986–1995	
Behavioral & Emotional Disorder	1986–1995	
Special Education	1991–1995	
Outcomes in Adolthood		
Education age 31	1973–1984	1966–1984
Employment/Earnings age 31	1973–1984	
Employment/Earnings age 26, 31, 36 & 41		1966–1971

The table summarizes the birth cohorts used for each outcome. See section 3 for more detail.

Table A2

Averages of Family Environment (Percent) by Cohort

Sample	Total Population			Sibling Sample		
Birth Cohorts	1966–1972	1973–1984	1985–1995	1966–1972	1973–1984	1985–1995
Background Information						
Child is male	51.19	51.19	51.34	51.27	51.20	51.48
Child's year of birth	1969	1978	1990	1969	1978	1990
# of Children in Family	2.63	2.51	2.59	2.78	2.66	2.77
Child's birth order	1.80	1.79	1.78	1.71	1.86	1.96
Mother immigrant	2.77	4.14	6.54	2.83	4.43	7.08
Father immigrant	2.56	4.37	7.26	2.67	4.52	7.70
Mother's age at birth (years)	25.34	26.30	27.97	24.44	26.28	28.20
Parental Education						
Mother <HS	50.75	39.76	28.24	48.83	38.64	27.74
Mother HS	33.66	35.56	40.71	34.38	35.99	40.29
Mother BA	15.59	24.68	31.05	16.79	25.37	31.97
Father <HS	36.58	28.85	24.08	35.46	27.74	23.07
Father HS	47.40	51.45	54.20	47.67	51.88	54.34
Father BA	16.02	19.71	21.72	16.87	20.38	22.59
Marital Status (Cohorts 1973-95) and Immigrant Background						
Married at own birth		72.29	54.32		77.69	62.30
Married at youngest sib birth		80.79	61.68		88.14	72.48
Family Structure at Youngest Sib Age 12						
Traditional family	79.89	71.78	67.73	82.00	77.79	73.43
Step-parent family	7.63	11.54	12.92	6.58	8.47	10.05
Single-parent family	12.48	16.67	19.35	11.42	13.74	16.52
N	424,597	661,024	628,518	301,270	531,497	446,864

The columns present averages by the two samples used for the analysis on educational attainment at age 31 and grade 9 completion on time (the total population and the sibling sample) for individuals born between 1966 and 1995. Note that the following variables are *not* reported as percent: year of birth, # of children in family, and birth order.

Table A3

Prenatal Inputs and Birth Outcomes

	Total Population OLS						Sibling Sample FE					
	Check-ups	Pregn. Comp.	Smoke	Pre-term	log(BW)	Low Apgar	Check-ups	Pregn. Comp.	Smoke	Pre-term	log(BW)	Low Apgar
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A												
Male	-0.09***	0.14***	-0.05	0.83***	3.37***	0.15***	-0.08***	0.09	0.33	0.60***	3.76***	0.10**
	(0.01)	(0.04)	(0.17)	(0.05)	(0.04)	(0.02)	(0.01)	(0.06)	(0.22)	(0.07)	(0.05)	(0.04)
Panel B												
Male	-0.08***	0.17	-0.07	0.94***	3.13***	0.23***	-0.08***	0.11	0.07	0.76***	3.70***	0.21*
	(0.02)	(0.11)	(0.46)	(0.13)	(0.11)	(0.06)	(0.03)	(0.17)	(0.65)	(0.21)	(0.15)	(0.11)
Male×Mom HS	-0.01	-0.01	-0.34	0.02	0.03	0.01	-0.02	-0.12	0.46	-0.05	0.07	0.08
	(0.02)	(0.11)	(0.45)	(0.12)	(0.10)	(0.06)	(0.02)	(0.15)	(0.59)	(0.18)	(0.13)	(0.10)
Male×Mom BA	0.00	-0.07	0.13	-0.32**	0.12	-0.02	0.01	-0.23	0.47	-0.34*	0.27*	0.01
	(0.02)	(0.12)	(0.49)	(0.13)	(0.11)	(0.06)	(0.03)	(0.17)	(0.66)	(0.19)	(0.15)	(0.10)
Male×Dad HS	-0.03*	0.10	-0.03	0.07	0.03	-0.04	0.02	0.12	-0.48	0.18	-0.21*	-0.14
	(0.02)	(0.11)	(0.44)	(0.12)	(0.10)	(0.06)	(0.02)	(0.16)	(0.58)	(0.18)	(0.13)	(0.10)
Male×Dad BA	-0.01	0.10	-0.11	0.16	0.06	0.01	-0.00	0.19	-0.81	0.25	-0.18	0.01
	(0.02)	(0.14)	(0.52)	(0.15)	(0.13)	(0.07)	(0.03)	(0.19)	(0.71)	(0.22)	(0.17)	(0.12)
Male×Married	0.01	-0.15*	0.33	-0.18*	0.30***	-0.11**	0.00	0.00	0.52	-0.25	0.15	-0.10
	(0.01)	(0.09)	(0.33)	(0.09)	(0.08)	(0.05)	(0.02)	(0.14)	(0.50)	(0.16)	(0.12)	(0.09)
N	874,850	874,850	291,211	855,825	853,686	865,474	564,826	564,826	95,936	549,545	547,422	554,777
Mean of Y	10.757	4.393	30.647	4.737	813.046	1.112	10.834	3.985	27.086	4.346	813.883	1.104
Prob > F	0.296	0.876	0.797	0.039	0.665	0.887	0.718	0.837	0.768	0.219	0.159	0.403

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. F-test of whether all the male-interaction terms are equal to zero. All estimates, except for those in Columns (1) and (7), are multiplied by 100. The samples consist of individuals born from 1980–1995. Estimates from each column in each panel come from separate regressions. *Check-ups* measures the total number of check-ups at general practitioner, midwife, and specialist during pregnancy. *Pregnancy complications* is a binary indicator taking the value 1 if the mother is diagnosed with preeclampsia, gestational hypertension, or gestational diabetes mellitus during pregnancy. *Smoke* indicates whether the mother smoked during pregnancy and is observed for birth cohorts 1991–1995. *Preterm* indicates whether the child was born before 37 weeks of gestation. *Log(BW)* represents the natural logarithm of birth weight. *Low Apgar* indicates whether the child had a five minutes Apgar score below 7, which is medically considered a low score. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), and a constant. All OLS models additionally control for family size dummies, parental immigration status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table A4

Labor Market Outcomes Across Ages (Birth Cohorts 1966–71)

	Employed				Earnings Percentile by Birth Cohort and Gender			
Age	26	31	36	41	26	31	36	41
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A								
Male	4.18***	4.30***	2.90***	1.38***	0.79***	0.58***	0.43**	0.05
	(0.18)	(0.17)	(0.18)	(0.19)	(0.18)	(0.18)	(0.18)	(0.18)
Panel B								
Male	6.89***	6.35***	4.32***	1.08	3.88***	1.13*	-1.16**	-2.16***
	(0.71)	(0.71)	(0.73)	(0.78)	(0.60)	(0.59)	(0.59)	(0.59)
Male×Mom HS	-2.30***	-2.29***	-1.41***	-1.56***	-4.16***	-1.41***	-0.42	-0.47
	(0.43)	(0.40)	(0.42)	(0.44)	(0.44)	(0.43)	(0.43)	(0.43)
Male×Mom BA	-1.53**	-2.10***	-1.19**	-1.38**	-4.86***	-1.46**	-0.82	-0.78
	(0.64)	(0.58)	(0.59)	(0.60)	(0.60)	(0.62)	(0.62)	(0.62)
Male×Dad HS	-0.95**	-0.76*	-0.41	-0.63	-1.95***	-0.93**	0.03	0.25
	(0.41)	(0.40)	(0.42)	(0.44)	(0.41)	(0.40)	(0.40)	(0.40)
Male×Dad BA	-1.13*	-0.47	-1.78***	-0.34	-6.39***	-2.61***	-0.58	-0.03
	(0.67)	(0.60)	(0.61)	(0.63)	(0.63)	(0.65)	(0.64)	(0.64)
Male×Trad 12	-1.41**	-0.72	-0.35	1.53**	0.87	1.15**	2.30***	2.77***
	(0.68)	(0.68)	(0.69)	(0.75)	(0.59)	(0.58)	(0.58)	(0.58)
Male×Step 12	-0.14	-0.31	-0.75	0.96	1.38	-0.45	0.32	1.25
	(1.08)	(1.07)	(1.13)	(1.19)	(0.92)	(0.92)	(0.92)	(0.91)
N	157,905	157,905	157,905	157,905	157,905	157,905	157,905	157,905
Mean of Y	91.89	92.83	92.17	91.08	53.22	53.61	53.33	52.99

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. The sample is the subsample of sibling sample with individuals born from 1966-71. Estimates in each column in each panel come from separate regressions. *Employed* takes the value one if the person has positive wage earnings or have employment as the main source of income including self-employment and zero otherwise. *Earnings percentile* measures the annual earnings percentile by gender and year of birth. All models control for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at the specified age, and a constant.

Table A5

Educational Attainment at Age 31 by Cohort: Male-Cohort-Family Environment Interactions

	Male (1)	Mom HS (2)	Mom BA (3)	Dad HS (4)	Dad BA (5)	Trad 12 (6)	Step 12 (7)
Panel A: Highest Completed Education (Months)							
Male×1966-70	-2.90***	-0.80**	-2.69***	2.80***	2.63***	0.62	-0.18
	(0.46)	(0.32)	(0.46)	(0.31)	(0.48)	(0.45)	(0.71)
Male×1971-74	-4.24***	-1.00***	-3.46***	1.25***	1.70***	1.32***	0.52
	(0.46)	(0.31)	(0.42)	(0.31)	(0.45)	(0.44)	(0.67)
Male×1975-78	-5.10***	-1.00***	-2.66***	0.92***	1.32***	0.74	-0.48
	(0.48)	(0.33)	(0.42)	(0.33)	(0.47)	(0.45)	(0.69)
Male×1979-84	-5.26***	-1.28***	-2.72***	0.64*	1.20**	-0.36	-0.32
	(0.52)	(0.35)	(0.42)	(0.36)	(0.49)	(0.48)	(0.71)
$Prob > F1$	0.00	0.79	0.46	0.00	0.14	0.07	0.74
$Prob > F2$	0.00	0.31	0.96	0.00	0.04	0.13	0.89
Panel B: At least HS Degree							
Male×1966-70	-2.44***	-2.31***	-2.60***	3.26***	2.29***	0.51	1.25
	(0.79)	(0.51)	(0.61)	(0.53)	(0.64)	(0.75)	(1.24)
Male×1971-74	-3.28***	-2.39***	-2.27***	1.19**	1.14*	2.43***	1.94*
	(0.77)	(0.49)	(0.56)	(0.52)	(0.59)	(0.70)	(1.12)
Male×1975-78	-2.31***	-1.10**	-1.98***	1.45***	0.49	0.94	-0.40
	(0.80)	(0.51)	(0.56)	(0.54)	(0.61)	(0.72)	(1.14)
Male×1979-84	-2.01**	-0.91*	-1.83***	1.78***	1.21*	-0.10	-1.49
	(0.84)	(0.53)	(0.56)	(0.57)	(0.64)	(0.74)	(1.14)
$Prob > F1$	0.70	0.07	0.80	0.02	0.23	0.07	0.13
$Prob > F2$	0.71	0.06	0.35	0.05	0.23	0.56	0.10
Panel C: BA Degree							
Male×1966-70	-6.09***	-1.12**	-6.64***	1.08**	3.05***	-2.70***	-0.88
	(0.63)	(0.52)	(0.82)	(0.47)	(0.85)	(0.64)	(0.98)
Male×1971-74	-7.78***	-2.52***	-9.30***	-0.17	3.67***	-2.81***	-1.06
	(0.65)	(0.52)	(0.75)	(0.50)	(0.80)	(0.65)	(0.97)
Male×1975-78	-11.24***	-2.33***	-6.66***	-1.36**	3.09***	-2.48***	-1.70
	(0.73)	(0.57)	(0.75)	(0.55)	(0.84)	(0.71)	(1.05)
Male×1979-84	-11.80***	-3.76***	-6.30***	-1.98***	3.41***	-3.87***	0.80
	(0.77)	(0.61)	(0.75)	(0.59)	(0.85)	(0.75)	(1.09)
$Prob > F1$	0.00	0.01	0.02	0.00	0.94	0.54	0.39
$Prob > F2$	0.00	0.00	0.76	0.00	0.76	0.24	0.25

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates for HS and BA degree are multiplied by 100. The sample is the sibling sample with individuals born from 1966–1984. $N = 755,850$. For each panel, all estimates come from one regression as specified in equation (3), i.e. a regression interacting the male-family environment interactions as well as the family environment variables with birth cohort dummies. The models control additionally for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant. The outcome in Panel A is highest completed education (in months) at age 31 with a mean of 160.80 months; the outcome in Panel B is *HS graduate or more*, indicating whether the person has at least 12 years of education by age 31 with a mean of 81.11 percent; the outcome in Panel C is *BA graduate*, indicating whether the person has at least a BA degree by age 31 with a mean of 31.54 percent. $Prob > F1$ reports the p-value from a joint F-test of whether all the estimates in the particular column are equal. $Prob > F2$ reports the p-value from a joint F-test of whether the estimates for the earliest and the last cohorts in the particular column are equal.

Table A6

Interactions between Mother's and Father's Education

	– Grade 9 –			– Age 31 –		
	On Time	GPA	Highest Edu	BA	Em-ployed	Earnings Pct.
	(1)	(2)	(3)	(4)	(5)	(6)
Male	-10.86***	-30.30***	-5.01***	-11.40***	5.59***	1.31***
	(0.60)	(1.38)	(0.39)	(0.59)	(0.55)	(0.41)
Male×Mom HS	0.70	-2.04	-0.95**	-3.73***	-2.18***	-1.50***
	(0.77)	(1.78)	(0.44)	(0.73)	(0.58)	(0.48)
Male×Mom BA	0.22	-1.44	-3.10***	-10.96***	-2.51***	-2.79***
	(0.98)	(2.32)	(0.65)	(1.14)	(0.81)	(0.70)
Male×Dad HS	0.48	-0.84	1.05***	-2.82***	-2.37***	-1.40***
	(0.71)	(1.66)	(0.35)	(0.54)	(0.47)	(0.37)
Male×Dad BA	0.61	3.41	-0.44	-1.84	-3.61***	-2.83***
	(1.32)	(3.01)	(0.77)	(1.33)	(0.99)	(0.83)
Male×Married	0.76**	-1.36*	0.23	-1.39**	0.29	1.27***
	(0.34)	(0.80)	(0.34)	(0.56)	(0.46)	(0.37)
Parental Education-Interactions:						
Male×Mom HS*Dad HS	0.03	0.94	-0.37	1.08	0.39	0.43
	(0.93)	(2.17)	(0.54)	(0.92)	(0.70)	(0.59)
Male×Mom HS*Dad BA	1.34	3.19	1.44	4.19**	2.31*	0.66
	(1.55)	(3.55)	(0.97)	(1.72)	(1.21)	(1.07)
Male×Mom BA*Dad HS	1.02	4.16	-0.55	3.35**	0.65	0.50
	(1.14)	(2.70)	(0.77)	(1.36)	(0.95)	(0.83)
Male×Mom BA*Dad BA	1.79	2.96	2.77***	9.72***	1.63	0.46
	(1.58)	(3.67)	(1.02)	(1.79)	(1.27)	(1.10)
N	335,241	288,667	355,090	355,090	365,676	365,676
Mean of Y	83.38	2.98	164.14	36.58	90.18	54.64
Prob > F	0.75	0.32	0.00	0.00	0.38	0.93

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates except for highest educational attainment are multiplied by 100. F-test of whether the additional Male×Mom Edu×Dad Edu interaction terms are jointly equal to zero. The sample consists of the sibling sample born from 1986–1995 for the two grade 9 outcomes and from 1973–1984 for the four age 31 outcomes. *Grade 9 on time* indicates whether the person completed grade 9 by age 16. *GPA* is an average of all grades given during grade 9 both from teacher assessment and final exams for all subjects and is standardized with mean zero and standard deviation of one by year of grade 9 completion for the total population. *Highest completed education* measures the length of highest completed education in months. *BA graduate* indicates whether the person has at least a BA degree. *Employed* takes the value one if the person has positive wage earnings or have employment as the main source of income including self-employment and zero otherwise. *Earnings percentile* measures the annual earnings percentile by gender and year of birth. All models control for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), a constant, and dummies for age at observation if not observed at age 31 for the adult outcomes.

Table A7

Interactions between Family Environment and Marital Status at birth

	– Grade 9 –		– Age 31 –			
	On Time	GPA	Highest Edu	BA	Em-ployed	Earnings Pct.
	(1)	(2)	(3)	(4)	(5)	(6)
Male	-10.84***	-31.03***	-3.73***	-8.94***	6.92***	1.65***
	(0.73)	(1.65)	(0.58)	(0.79)	(0.89)	(0.61)
Male×Mom HS	0.94	-1.09	-2.64***	-4.62***	-4.53***	-2.55***
	(0.76)	(1.75)	(0.81)	(1.31)	(1.11)	(0.87)
Male×Mom BA	1.46*	2.11	-4.20***	-9.30***	-3.20***	-2.59***
	(0.83)	(1.96)	(0.94)	(1.57)	(1.21)	(0.98)
Male×Dad HS	-0.18	0.04	0.16	-5.24***	-2.90***	-1.64**
	(0.74)	(1.69)	(0.74)	(1.13)	(1.05)	(0.78)
Male×Dad BA	2.12**	3.95*	0.52	0.21	-3.10**	-2.08*
	(0.94)	(2.19)	(1.10)	(1.86)	(1.41)	(1.16)
Male×Married	0.46	-1.48	-1.25**	-5.10***	-1.49	0.72
	(0.89)	(2.05)	(0.63)	(0.87)	(0.95)	(0.66)
Marital Status at Birth-Interactions:						
Male×Mom HS*Married	-0.23	-0.19	1.71**	1.89	3.11***	1.50*
	(0.90)	(2.10)	(0.85)	(1.39)	(1.15)	(0.91)
Male×Mom BA*Married	-0.59	-1.15	1.60	2.64	1.37	0.12
	(0.98)	(2.33)	(0.99)	(1.67)	(1.27)	(1.04)
Male×Dad HS*Married	1.21	0.52	0.74	3.63***	0.85	0.51
	(0.89)	(2.05)	(0.79)	(1.21)	(1.10)	(0.83)
Male×Dad BA*Married	-0.43	1.37	1.16	3.69*	1.00	-0.50
	(1.10)	(2.59)	(1.16)	(1.97)	(1.48)	(1.23)
N	335,241	288,667	355,090	355,090	365,676	365,676
Mean of Y	83.38	2.98	164.14	36.58	90.18	54.64
Prob > F	0.21	0.98	0.04	0.00	0.06	0.33

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates except for highest educational attainment are multiplied by 100. F-test of whether the additional Male×Family Environment×Married interaction terms are jointly equal to zero. The sample consists of the sibling sample born from 1986–1995 for the two grade 9 outcomes and from 1973–1984 for the four age 31 outcomes. *Grade 9 on time* indicates whether the person completed grade 9 by age 16. *GPA* is an average of all grades given during grade 9 both from teacher assessment and final exams for all subjects and is standardized with mean zero and standard deviation of one by year of grade 9 completion for the total population. *Highest completed education* measures the length of highest completed education in months. *BA graduate* indicates whether the person has at least a BA degree. *Employed* takes the value one if the person has positive wage earnings or have employment as the main source of income including self-employment and zero otherwise. *Earnings percentile* measures the annual earnings percentile by gender and year of birth. All models control for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), a constant, and dummies for age at observation if not observed at age 31 for the adult outcomes.

Table A8

Robustness of Family Structure: Education and Labor Market Outcomes at Age 31

	Compl Edu		BA		Employed	
	(1)	(2)	(3)	(4)	(5)	(6)
Male	-4.57***	-4.63***	-10.13***	-10.34***	6.01***	5.32***
	(0.40)	(0.47)	(0.60)	(0.71)	(0.56)	(0.65)
Male×Mom HS	-1.11***	-1.05***	-2.88***	-2.66***	-1.76***	-1.65***
	(0.25)	(0.26)	(0.43)	(0.46)	(0.30)	(0.32)
Male×Mom BA	-2.83***	-2.77***	-6.99***	-6.88***	-1.95***	-1.78***
	(0.30)	(0.31)	(0.54)	(0.57)	(0.36)	(0.38)
Male×Dad HS	0.80***	0.61**	-1.99***	-2.09***	-2.16***	-2.09***
	(0.25)	(0.26)	(0.42)	(0.45)	(0.32)	(0.33)
Male×Dad BA	1.53***	1.49***	3.55***	3.97***	-2.27***	-2.22***
	(0.34)	(0.36)	(0.61)	(0.65)	(0.42)	(0.44)
Male×Trad 12	-0.20		-3.83***		-0.40	
Oldest	(0.38)		(0.58)		(0.51)	
Male×Step 12	-0.87		-1.52*		0.05	
Oldest	(0.58)		(0.87)		(0.79)	
Male×Trad 12		0.00		-3.93***		0.14
		(0.45)		(0.69)		(0.61)
Male×Step 12		-0.93		-2.05**		0.80
		(0.68)		(1.04)		(0.93)
N	355,090	317,649	355,090	317,649	365,676	327,055
Mean of Y	164.14	165.05	36.58	37.52	90.18	90.65
Sample:						
All Siblings	X		X		X	
Only Same FS12		X		X		X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates except for highest educational attainment are multiplied by 100. The sample consists of the sibling sample born from 1973–1984. The *All Siblings* sample tests the robustness of the main results by using family structure at age 12 of the oldest child instead of the youngest. The sample of *Only Same FS12* tests the robustness of the main results by only using the sample of families in which children experience the same family structure at age 12. *Highest completed education* measures the length of highest completed education in months. *BA graduate* indicates whether the person has at least a BA degree. *Employed* takes the value one if the person has positive wage earnings or have employment as the main source of income including self-employment and zero otherwise. *Earnings percentile* measures the annual earnings percentile by gender and year of birth. All models control for family fixed effects, year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant.

Table A9

Total Population vs Siblings: Grade 9

Dependent Var.	Grade 9 On Time						GPA					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A												
Male	-9.30***	-9.82***	-8.86***	-8.71***	-8.97***	-9.07***	-29.64***	-30.05***	-30.93***	-31.57***	-30.41***	-30.76***
	(0.09)	(0.35)	(0.12)	(0.19)	(0.15)	(0.15)	(0.26)	(0.98)	(0.35)	(0.66)	(0.37)	(0.36)
Panel B												
Male	-11.42***	-10.88***	-11.28***	-11.61***	-11.05***	-11.03***	-27.07***	-25.50***	-29.31***	-28.63***	-29.86***	-31.11***
	(0.26)	(0.84)	(0.38)	(0.61)	(0.47)	(0.48)	(0.67)	(2.33)	(1.01)	(1.82)	(1.12)	(1.10)
Male×Mom HS	0.56**	-0.55	0.89***	0.78	0.94**	0.80*	-3.00***	-4.72**	-1.87**	-1.68	-1.87*	-1.20
	(0.24)	(0.86)	(0.32)	(0.54)	(0.40)	(0.41)	(0.64)	(2.31)	(0.90)	(1.65)	(0.99)	(0.96)
Male×Mom BA	0.83***	0.26	0.80**	0.32	1.20***	1.07**	-2.26***	-6.27**	-0.43	-1.45	0.45	1.28
	(0.27)	(0.98)	(0.35)	(0.58)	(0.43)	(0.45)	(0.72)	(2.62)	(0.98)	(1.81)	(1.09)	(1.06)
Male×Dad HS	0.80***	0.28	0.79**	1.00*	0.65	0.65	-1.98***	-1.34	-1.66*	-3.24**	-0.45	0.39
	(0.24)	(0.83)	(0.32)	(0.54)	(0.40)	(0.41)	(0.64)	(2.24)	(0.90)	(1.65)	(0.98)	(0.96)
Male×Dad BA	2.53***	2.84***	2.45***	3.35***	1.75***	1.71***	2.32***	4.53	2.53**	0.90	3.87***	4.95***
	(0.29)	(1.10)	(0.39)	(0.63)	(0.48)	(0.49)	(0.80)	(2.92)	(1.08)	(1.98)	(1.20)	(1.16)
Male×Married	1.22***	1.44**	1.13***	1.66***	0.77**	0.76**	-1.65***	-3.19*	-1.40*	-2.04	-1.06	-1.35*
	(0.18)	(0.72)	(0.26)	(0.42)	(0.34)	(0.34)	(0.50)	(1.90)	(0.74)	(1.31)	(0.82)	(0.80)
N	579,049	42,836	335,241	143,083	192,158	335,241	524,330	38,431	288,667	123,252	165,415	288,667
Mean of Y	82.45	80.59	83.38	83.48	83.31	83.38	-1.41	4.18	2.98	4.63	1.74	2.98
Sample:												
Total Population	X						X					
One-Child Families		X						X				
Sibling Sample:												
All Sibs			X			X			X			X
Same Sex Sibs				X						X		
Mixed Sex Sibs					X						X	
Estimation:												
OLS	X	X	X	X	X		X	X	X	X	X	
FE						X						X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates are multiplied by 100. All samples consist of children born from 1986–1995. Estimates from each column in each panel come from separate regressions. *Grade 9 on time* indicates whether the person completed grade 9 by age 16. *GPA* is an average of all grades given during grade 9 both from teacher assessment and final exams for all subjects and is standardized with mean zero and standard deviation of 1 by year of grade 9 completion for the total population. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), and a constant. All OLS models additionally control for family size dummies, parental immigrant status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Table A10

Total Population vs Siblings: Age 31

Dependent Var.	Highest Compl Edu (Months)						Employed					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Panel A												
Male	-5.37***	-5.78***	-5.37***	-5.51***	-5.25***	-5.20***	2.86***	1.84***	2.79***	2.55***	2.96***	2.95***
	(0.07)	(0.24)	(0.09)	(0.16)	(0.11)	(0.10)	(0.08)	(0.27)	(0.10)	(0.15)	(0.13)	(0.13)
Panel B												
Male	-5.08***	-4.63***	-5.16***	-5.62***	-4.78***	-4.98***	4.79***	2.75***	4.76***	3.58***	5.64***	5.49***
	(0.17)	(0.53)	(0.30)	(0.52)	(0.36)	(0.36)	(0.22)	(0.63)	(0.39)	(0.61)	(0.50)	(0.51)
Male×Mom HS	-1.29***	-1.31**	-1.36***	-1.47***	-1.27***	-1.11***	-1.48***	-0.40	-1.70***	-1.60***	-1.76***	-1.80***
	(0.15)	(0.54)	(0.21)	(0.35)	(0.25)	(0.25)	(0.18)	(0.64)	(0.23)	(0.36)	(0.30)	(0.30)
Male×Mom BA	-3.00***	-2.87***	-2.98***	-2.94***	-3.00***	-2.82***	-1.75***	-1.24	-1.80***	-1.66***	-1.88***	-1.96***
	(0.19)	(0.65)	(0.25)	(0.44)	(0.30)	(0.30)	(0.22)	(0.77)	(0.28)	(0.44)	(0.35)	(0.36)
Male×Dad HS	0.34**	-0.51	0.71***	0.68*	0.74***	0.79***	-1.39***	0.32	-1.50***	-0.63	-2.11***	-2.19***
	(0.16)	(0.54)	(0.21)	(0.37)	(0.25)	(0.25)	(0.19)	(0.64)	(0.24)	(0.39)	(0.31)	(0.32)
Male×Dad BA	1.05***	0.07	1.24***	0.79	1.56***	1.53***	-2.08***	-1.81**	-2.28***	-2.21***	-2.30***	-2.30***
	(0.22)	(0.73)	(0.29)	(0.50)	(0.35)	(0.34)	(0.25)	(0.86)	(0.32)	(0.51)	(0.41)	(0.42)
Male×Married	0.66***	0.46	0.44	0.92*	0.03	0.21	0.21	-0.42	0.40	0.89	-0.02	0.25
	(0.15)	(0.46)	(0.28)	(0.48)	(0.34)	(0.34)	(0.18)	(0.55)	(0.35)	(0.54)	(0.45)	(0.46)
N	632,508	54,627	355,090	148,725	206,365	355,090	643,219	55,619	365,676	153,201	212,475	365,676
Mean of Y	162.97	163.61	164.14	164.54	163.86	164.14	89.54	88.09	90.18	90.67	89.84	90.18
Sample:												
Total Population	X						X					
One-Child Families		X						X				
Sibling Sample:												
All Sibs			X			X			X			X
Same Sex Sibs				X						X		
Mixed Sex Sibs					X						X	
Estimation:												
OLS	X	X	X	X	X		X	X	X	X	X	
FE						X						X

Standard errors in parentheses, clustered at the family level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. All estimates for employment are multiplied by 100. All samples consist of children born from 1973–1984. Estimates from each column in each panel come from separate regressions. *Highest completed education* measures the length of highest completed education in months. *Employed* takes the value one if the person has positive wage earnings or have employment as the main source of income including self-employment and zero otherwise. All models control for year of birth dummies, month of birth dummies, birth order dummies, maternal age at birth (linear, squared, and cubed), dummies for age at observation if not observed at age 31, and a constant. All OLS models additionally control for family size dummies, parental immigrant status, and those variables of family environment that are interacted with the male dummy, and the FE models control for family fixed effects.

Chapter 4 – The Intergenerational Transmission of Time Preferences Persists Across Four Decades

The Intergenerational Transmission of Time Preferences Persists Across Four Decades*

Anne Ardila Brenøe[a,b] Thomas Epper[a,c,d]

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Abstract

We study the intergenerational transmission of time preferences, using an experimentally validated survey measure. Parents' and children's impatience is measured four decades apart, thereby eliminating concerns regarding reverse causality. Our results show a substantial transmission of impatience from parents to children. This correlation is insensitive to the inclusion of comprehensive sets of administratively reported controls. We further show that mothers differentially and more strongly transmit impatience to daughters relative to sons, even when comparing siblings. This suggests that nurture affects children's impatience. Finally, the strength of the transmission does not diminish as children age, emphasizing the persistence of preference propagation.

JEL classification: D1, J2, J6

Keywords: Intergenerational transmission of preferences, impatience, time preferences.

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1 Introduction

Patient people generally experience better lifetime outcomes than their more impatient counterparts.[1] Measures of time preferences elicited during childhood are predictive for how individuals fare in later life, for instance, in terms of education, health, and earnings (Golsteyn, Grönqvist and Lindahl, 2014; Mischel, Shoda and Peake, 1988). If parents transmit their time preferences to children, it may help explain an important part of the cross-generational correlation of outcomes—such as the position in the wealth distribution, health status, and educational attainment. Getting deeper insights in how impatience propagates across generations may help policy makers develop programs counteracting potentially detrimental behaviors sustained across generations, like underinvestment in human or health capital, notorious undersaving, and excessive credit card borrowing. Yet, we do not know much about the origins of time preferences and the transmission of such preferences across generations. Are time preferences passed on to the next generation? Is such transmission permanent or only short-lived? What are channels through which preferences carry over from parents to offspring? In this paper, we address these questions to improve our understanding of the intergenerational propagation of time preferences.

We study the transmission of impatience from parents to children, using data from a unique Danish survey linked to high-quality administrative data. Parents and children answer the same time preference question four decades apart, eliminating concerns regarding reverse causality. In our central question, respondents picked one out of three possible income profiles. Our survey measure is internally and externally valid: First, we show that the individuals whom we categorize as being impatient have significantly worse socio-economic outcomes in adulthood, controlling for a wide range of childhood family characteristics. Second, Epper et al. (2018) document that our survey question strongly correlates with time preferences elicited in a real-incentivized experiment in a large and heterogeneous population.

Our results show a substantial transmission of impatience from parents to children. This correlation is insensitive to the inclusion of a wide array of controls, including grandparental and parental socio-economic status and child endowments.

[1] See e.g. Ayduk et al. (2000); Chabris et al. (2008); Epper et al. (2018); Golsteyn, Grönqvist and Lindahl (2014); Meier and Sprenger (2012); Mischel, Shoda and Peake (1988); Shoda, Mischel and Peake (1990); Sutter et al. (2013).

The association between mothers' and daughters' time preferences is particularly strong, while mothers' and sons' preferences are uncorrelated. We further show that mothers' differential transmission of impatience to daughters relative to sons remains similar in magnitude when comparing sisters to their brothers. From such comparison, we can exclude family socio-economic status and genetic transmission as major mediators of the intergenerational propagation of preferences from mothers to daughters. This suggests that role-modeling is an important mechanism for the transmission of impatience in mother-daughter dyads. This is consistent with the finding in Brenøe and Lundberg (2017) that mothers affect daughters more positively than sons in terms of educational attainment. At the same time, we do not find that fathers' transmission of impatience differs by child gender. Finally, we find evidence of persistent effects, as the correlation between parents' and children's preferences does not diminish when children age. The transmission from fathers to children is indeed stronger for older relative to younger children.

The literature on intergenerational transmission of time and other related economic preferences is small but recently expanding. Overall, this literature finds that risk, time, and social preferences between parents and their children correlate (Alan et al., 2017; Bartling et al., 2010; Brown and van der Pol, 2015; Chowdhury, Sutter and Zimmermann, 2018; Dohmen et al., 2012; Gauly, 2016; Kosse and Pfeiffer, 2013). One major empirical limitation in the existing studies is, however, the very short time delay between the elicitation of parents' and children's preferences. Most studies measure the preferences contemporaneously, while only a few studies feature a short time delay (see Appendix Table A1 summarizing the literature on the intergenerational transmission of time preferences). Because children may affect parents, correlations between parents' and children's preferences do not necessarily provide evidence of an intergenerational transmission from parents to children. Consistent with our findings, some studies find that mothers only transmit their preferences to daughters, but not to sons,[2] while others either do not document or do not find gender differences in the transmission process.[3] Moreover, similar to our findings, these studies do not generally find that the correlations weaken when controlling for socio-economic status. This suggests that

[2]Alan et al. (2017) study the transmission of risk attitudes from mothers to children (age 7–8) and find strong correlations between mothers' and daughters', but not between mothers' and sons', preferences. Gauly (2016) find a similar pattern for patience.

[3]Dohmen et al. (2012) study correlations between parents' and their adult children's risk and trust preferences and do not find differences by parental gender.

the socio-economic childhood environment does not mediate the intergenerational transmission of preferences.

Another related strand of the literature examines how childhood socio-economic status correlates with children's economic preferences. Findings within this literature suggest that children from high relative to low socio-economic backgrounds, for example, exhibit a substantial difference in economic and social preferences and willingness to compete (Almås et al., 2015; Bauer, Chytilová and Pertold-Gebicka, 2014; Deckers et al., 2017). Deckers et al. (2017), for instance, conclude that socio-economic status is a powerful predictor of children's preferences and that differences in parental investment represent an important mediator of the relationship. Common to these studies is, however, that they do not show how the gap in preferences by socio-economic status changes once controlling for parental preferences. Thus, it is not clear whether the correlation between family resources and children's preferences is because the socio-economic environment influences preferences or because parents transmit preferences to children. Parents with preferences associated with worse socio-economic outcomes may, for instance, transmit their potentially unfavorable preferences to children independent of the socio-economic environment. Disentangling the mechanisms behind children's preference formation is highly relevant for policy purpose, as such knowledge is key for the development of successful interventions aiming at increasing social mobility in disadvantaged families.

While we are not the first to study the intergenerational transmission of preferences, we contribute to the literature in four important ways. First, in contrast to existing studies, we measure parents' preferences before they have children, thereby ruling out reverse causality concerns. This data feature further permits conclusions on the persistence of propagation effects. Second, we consider a large, representative sample with exceptionally low attrition rates in a high-income country. While existing work typically studies more selected and homogeneous samples, little is known about how this evidence extends to a representative Western society. Third, we have linked the survey data to comprehensive administrative registries to obtain precise measures of demographic characteristics, parental fertility, and parental socio-economic status during children's childhood for both parents (not only for the parent in the survey). This enables a rigorous investigation of whether socio-economic status during childhood mediates the transmission of impatience.

Fourth, for the majority of parents, we observe multiple children, allowing us to explore whether the transmission differs by gender and age of children, while keeping the family environment constant. Thereby, we can say more about transmission channels than what is normally possible in this type of study.

2 Data

Our main data source is the Danish Longitudinal Survey of Youth-Children (DLSY-C)[4], which we link to high-quality administrative data on the entire Danish population from 1980 through 2016. This combined data set provides unique possibilities for studying the intergenerational transmission of time preferences. The DLSY is a longitudinal study of 3,151 individuals born around 1954, whom we will refer to as *parents*. In 1968, these original respondents attended 152 different seventh grade school classes that were sampled to be nationally representative. The parents have subsequently been interviewed throughout their adult life with high response rates; around 75 percent of the original individuals participated in the last wave in 2004. In addition, the parents of the respondents (henceforth referred to as *grandparents*) were interviewed in 1969, making it possible to control for the parents' socio-economic environment during their childhood. Finally, all *children* born to the DLSY respondents and at least 14 years old were interviewed in 2010, with an extraordinarily high response rate of 81 percent. We therefore have information on three generations: grandparents, parents, and children.

In 1973 (at age 19), the parents answered a question regarding their time preferences. The question is:

If you were offered three jobs now and you should choose, which one would you take?

(a) *a job with average pay right from the beginning*,

(b) *a job with low pay the first two years, but high pay later*, or

(c) *a job with very low pay the first four years, but very high pay later*.

In what follows, we categorize respondents answering (a) as impatient. The children answered the very same question nearly four decades later in 2010 when they

[4]The data set is provided by VIVE (The Danish Center for Social Science Research) and consists of the Danish Longitudinal Survey of Youth (DLSY) and DLSY-C.

were 27 years on average, with ages ranging between 14 and 40 years.[5] The timing of the parents' elicitation of time preferences allows us to rule out any issue of potential reverse causality, as only 2.8 percent ($N = 87$) of the children were born by 1973 and only 10 children were more than one year. As a robustness check, however, we exclude children born at the time of the parents' response to the time preference question and reach similar results. Thus, our empirical setup gives us the power to study intergenerational transmission of impatience in the absence of reverse causality concerns.

We observe time preferences for 3,101 children and 1,829 parents.[6] Table 1 presents descriptive statistics for the full sample. While 35.2 percent of children are labeled as impatient, 25.8 percent of parents are so. Considering gender differences in impatience, 26.1 percent of mothers, 24.6 percent of fathers, 38.1 percent of daughters, and 32.1 percent of sons are categorized as impatient.[7] The sample is balanced with respect to child and parent gender. On average, children have 1.5 siblings, while parents to the children in the sample have 2.2 children by 2016 and have 1.7 children in the sample (54.7 percent have at least two children in the sample). Seventy-three percent of children lived with both biological parents at age 16 and their parents have, on average, completed around 13 years of education.

As already mentioned, we further link the survey data to rich administrative data, including the Medical Birth Registry and several separate registers on education, income, (un)employment, fertility, and family structure. Therefore, in addition to the ample information on grandparents' socio-economic status during parents' childhood observed in the DLSY, we observe the socio-economic status experienced by the children during their childhood. Although we only observe impatience for one of the parents, we do observe both parents in the registers. Thus, we observe parents' complete fertility history, labor market experience, and educational

[5] Due to data protection rules, we are unable to report the exact maximum age.

[6] We do not observe all the original 3,151 DLSY respondents in the sample of parents: 618 individuals did not have any children by 1996; of those with at least one child by 1996, 301 individuals did not have an impatience observation; of those with at least one child by 1996 and with an impatience observation, 390 individuals did not have a child surveyed in 2010.

[7] Experiments confronting subjects (usually students) with smaller sooner versus larger later rewards typically find that females are more patient than males (see e.g. Dittrich and Leipold (2014)). We suspect that this opposing finding is due to contextuality and the samples under consideration. The survey question we use explicitly asks about choosing a wage profile, while experimental measures typically involve rather abstract allocations of monetary amounts. Older children (those in their 30s) in our data tend to be more impatient [Appendix Figure A1].

Table 1
Descriptive Statistics

	Mean	Std.Err.
	(1)	(2)
Panel A: Child		
Child Impatient	0.352	0.009
Daughter Impatient	0.381	0.012
Son Impatient	0.321	0.012
Parent Impatient	0.258	0.008
Mother Impatient	0.271	0.011
Father Impatient	0.245	0.011
Daughter	0.519	0.009
Mother	0.521	0.009
Child Age (years)	27.092	0.101
Twin	0.019	0.002
Birth Weight (grams)	3427	10.111
Lives with both parents at age 16	0.734	0.008
Mother's Years of Education	12.672	0.047
Father's Years of Education	13.051	0.051
# of Siblings	1.469	0.017
# of Siblings in sample	1.042	0.016
Observations	3,101	
Panel B: Parent		
Mother Impatient	0.261	0.014
Father Impatient	0.246	0.015
# of Children	2.225	0.020
# of Children in sample	1.695	0.018
Observations	1,829	

This table presents descriptive statistics for the sample of children and their parents. **Panel A** contains means and standard errors for all children with a measure on their own and their parent's time preferences. **Panel B** contains the respective information for all parents of the children in Panel A. Note that we observe one parent per family only, i.e. either the mother or the father. Differences between Panel A and B can be explained by the fact that parents may have multiple children.

attainment.[8] For the children, we observe their date of birth, birth outcomes, and educational attainment by 2016.

In comparison to experimental measures of time preferences (see Frederick, Loewenstein and O'Donoghue (2002) for a review), our survey measure has both advantages and disadvantages. The possibly most important advantage is that our survey question is short, simple, and less abstract than typical experimental allocation choices. Specifically, our question asks subjects about their choice in a real-life situation with substantial economic consequences. This contrasts experimental measures, typically asking subjects to repeatedly choose between comparatively small sooner amounts and slightly larger later amounts (usually materializing within some weeks or a few months). This context-dependence might also be viewed as a shortcoming of our measure, in that considerations other than pure time preferences might lead subjects to choose a particular wage profile. Risk averse individuals may, for instance, choose the average pay fearing they would not reach the high pay (although the question does not explicitly associate risk with future pay rises). In our sample, children's time and risk preferences do indeed correlate. This is consistent with the broader literature, which demonstrates that carefully designed experiments tailored to elicit time preferences are also prone to such confounds.[9] Despite this, our results on impatience remain robust once we control for risk preferences.

Importantly, Epper et al. (2018) document that the survey measure we use highly correlates with experimental measures of time preferences. More precisely, Epper et al. (2018) validate the DLSY impatience measure both internally and externally. First, in a large-scale online experiment with 4,152 Danes from the broad population born between 1967 and 1986, they demonstrate that men's and women's answers to the DLSY question are highly correlated with a preference measure inferred from an experiment with real-monetary incentives. Second, for our sample of parents, they show that the subjects we classify as impatient have a consistently lower percentile rank in the within-cohort wealth distribution over a 15-year period. Moreover, we illustrate in Appendix Table A2 that impatient parents experience substantially worse outcomes, both in terms of educational attainment and labor

[8]For the labor market outcomes, we restrict the focus to the years 1980–2004 (i.e. through age 50 of the parents) to proxy children's childhood family environment (the average child turned 21 years in 2004).

[9]See Epper and Fehr-Duda (2015) for a discussion. Unfortunately, we do not observe parents' risk preferences.

market performance. Impatient mothers (fathers), for instance, have 0.46 (0.60) fewer years of education and earn 44 (26) log-points less during age 26 through 50 than patient mothers (fathers).[10] Similarly, impatient daughters (sons) have attained 0.50 (0.52) fewer years of education by 2016. These findings demonstrate that the DSLY measure captures impatience well and that it is a good predictor of real-life economic outcomes.

Nevertheless, a final concern could be that—especially—women wishing to have children early might choose the impatient option, not because they per se are impatient but because they want a stable income to be able to afford having children in the near future. Thus, a stronger correlation between mothers' and daughters' time preferences than between mothers' and sons' could be due to a stronger transmission of fertility preferences between mothers and daughters. As a robustness check, proxying parents' preferred fertility with their revealed fertility,[11] we only consider children whose surveyed parent had their first child after 1977, i.e. when the wage increase in the most patient wage profile would be implemented. Thereby, we exclude parents who might have answered the time preference question considering their future fertility plans; the results remain similar. Appendix Table A4 further explores associations between parents' fertility preferences at age 22 and their realized lifetime fertility. Women's fertility preferences are independent of their time preferences, while impatient women are more likely to have children early. In contrast, impatient men are more likely, at age 22, to desire not having any children and are accordingly less likely to have any (recognized) children by age 62.

3 Empirical Strategy

Our empirical analysis proceeds in three steps. First, we examine the correlations between parents' and children's preferences, while extensively testing the robustness of the transmission by adding a wide set of controls. Second, we investigate whether the transmission of impatience for the full sample is driven by specific

[10] Note, these differences in economic outcomes by impatience category cannot be explained by justification bias, as parents answer the time preference question before their outcomes are measured.

[11] We acknowledge that this is an imperfect proxy for people's actual fertility preferences at the time of the elicitation of time preferences.

child-parent gender dyads and whether childhood socio-economic status mediates some of the impatience propagation. Third, we compare siblings to study whether parents differentially transmit their impatience to daughters relative to sons and older relative to younger children, while keeping the family environment and genetic factors fixed.

The first part of the analysis studies the conditional correlations between parents' and children's preferences. For this, we specify the following linear probability model for the full sample:[12]

$$\begin{aligned}\text{Impatient}_c = \alpha_1 \text{Impatient}_p + \alpha_2 \text{Impatient}_p \times \text{Mother}_p + \\ \alpha_3 \text{Mother}_p + Q_c' \zeta + R_p' \delta + X_g' \epsilon + \theta_s + \nu_{cp},\end{aligned} \tag{1}$$

where c denotes the child, p the parent, and g the grandparent. *Impatient* indicates whether the individual is impatient (1) or not (0). To allow for a differential transmission from mothers relative to fathers, we include an interaction term between parental impatience and gender of the parent. ν_{cp} denotes the error term; we cluster the standard errors at the parent level to allow for serial correlation in the outcome between siblings.

To shed light on the nature of the intergenerational transmission of impatience, we examine these correlations while stepwise adding extensive vectors of background characteristics. First, we add a vector of child demographic characteristics, Q,[13] that adjusts for potential differences in child impatience due to age and gender, among others. Second, we add a vector of parental demographic characteristics, R.[14] Third, we include school fixed effects for the surveyed parent's school in 1968, θ_s, as this was the original level of sampling. Fourth, to control for differences in parents' socio-economic status during their childhood, we add a vector of grandparents' socio-economic characteristics, X.[15] Because parental impatience correlates

[12] The results are qualitatively robust to non-linear specification (not reported).

[13] This vector of *child demographics* includes indicators for being female, five-year age intervals, birth order, and being twin.

[14] This vector of *parent demographics* includes indicators for being born before 1954, born after 1954, child-parent gender combination, and birth order.

[15] This vector of *grandparent SES* includes grandparental attitudes towards child education and work; an index for the grandparents' educational investment in the parent; quadratic taxable income in 1967 reported by the tax authorities; quadratic number of grandparents' children; indicators for the grandmaternal/paternal level of education, vocational training/education, grandmother/grandfather has subordinates, grandmother is housewife, gender of the surveyed grandparent, the parent lives with both parents at age 14, and indicators for missing observations for

with parents' adult (and children's childhood) socio-economic status and because the latter may be an important mediator of the intergenerational transmission, we prefer not to control for such variables in this part of the analysis. However, as a robustness check, we finally include parental adult socio-economic controls[16] and child endowments.[17]

After examining the sensitivity of the correlations when adding the different vectors of controls, the second part of the analysis studies more rigorously whether and how the intergenerational transmission differs by child and parent gender and the role of childhood socio-economic status. For this, we always estimate the model separately for each child-parent gender combination. The general model is:

$$\begin{aligned}\text{Impatient}_c = \beta_1 \text{Impatient}_p + \beta_2 \text{Impatient}_p \times \text{SES}_p + \\ \beta_3 \text{SES}_p + Q_c' \zeta + R_p' \delta + X_g' \epsilon + \theta_s + \nu_{cp},\end{aligned} \tag{2}$$

where SES_p denotes the socio-economic family environment during the child's childhood, i.e. *both* parents' adult socio-economic status.[18] To test whether the transmission of impatience is dominated by specific child-parent gender dyads, we first estimate Equation (2) by omitting the SES_p components. Next, we omit the Impatient_p components to check whether we can replicate the finding of an SES gap in children's time preferences as reported by Deckers et al. (2017). Finally, we estimate the full equation to see whether the inclusion of the childhood socio-economic environment changes the correlation between parents' and children's impatience.

The third part of the analysis investigates whether parents differentially transmit impatience to daughters relative to sons and whether the transmission of impatience persists as children age, while keeping the childhood family environment and genetic factors constant. We do this by including sibling fixed effects, μ_p, and

the different control variables.

[16]This vector of *parent SES* includes the surveyed parent's spatial, verbal, and inductive abilities measured at age 14; the mother's and father's length of education, cumulated work experience through 2004, cumulated length of unemployment through 2004, the natural logarithm of average annual labor earnings 1980–2004; quadratic number of children; indicators for the child lives with both parents at age 16, the mother/father has children with another person than the parent, and missing observations for the different control variables.

[17]This vector of *child endowments* includes squared birth weight, child IQ, standardized length of highest completed education by 2016 by cohort; indicators for being born preterm and missing observations for the different control variables.

[18]Empirically, we construct an *SES Index* (standardized with mean zero and standard deviation one), using the first principal component from a principal component analysis; see Appendix Table A3.

eliminate thereby potential time-invariant characteristics within the same sibship. We estimate this model for the sample of mothers and fathers, separately:

$$\text{Impatient}_c = \gamma_1 \text{Impatient}_p \times C_c + Q_c'\zeta + \mu_p + \nu_{cp}, \tag{3}$$

where C is the child's characteristic of interest (gender and age). As long as we do not omit any important time-varying variables in this model, we can interpret γ_1 as the causal effect of parental impatience on the gender (age) gap in child impatience.[19] This strategy thus tests the robustness of the second strategy's results on gender differences in the transmission of preferences. Moreover, although age at first birth is endogenous (and therefore examining heterogeneity in the transmission by child age for the full sample does not make much sense), we can say whether the transmission persists or fades out as children age in this setup. For each parent, we compare siblings born earlier versus later and are therefore older versus younger at the time of the interview, keeping the parent's age at first birth constant. Consequently, this approach comparing siblings of the same parent but of different gender and age provides a fruitful setting for getting closer to mechanisms behind the intergenerational transmission of preferences. It helps investigate whether the intergenerational transmission of impatience is due to differences in the social environment (gender) rather than genetics (accounted for by the inclusion of sibling fixed effects) and to study the persistence (age) of the transmission.

4 Results

We present the results in the order introduced in Section 3. First, Table 2 provides an overview of correlations between parents' and children's time preferences. Second, Table 3 examines the transmission of impatience across mother/father and daughter/son pairs and the mediating role of socio-economic status. Finally, Table 4 further tests the robustness of the results on gender differences in the transmission process and studies the persistence of the effects, including siblings fixed effects.

[19]See Brenøe and Lundberg (2017) for a more elaborate discussion of this empirical strategy.

4.1 Intergenerational Transmission of Impatience

The main finding in Table 2 is that parents significantly transmit impatience to their children and that this transmission is robust to the inclusion of comprehensive sets of controls. The raw correlation between parents' and children's preferences suggests that children with an impatient parent are 6.3 percentage points more likely to be impatient [Column (1)], corresponding to an increase of 18 percent relative to the mean. Moreover, we cannot reject that the strength of the transmission is independent of the parent's gender. Column (2) further reveals that daughters are around 6 percentage points more likely to be impatient than sons; a gender difference that is stable across all models in Table 2.

It is noteworthy that the magnitude of the intergenerational transmission remains similar when adding child and parental demographic variables and parents' school fixed effects [Column (3)]. Neither does the inclusion of a rich set of grandparental socio-economic status alter the estimated intergenerational transmission [Column (4), representing our preferred model]. Finally, Column (5) shows that the correlation between parents' and children's impatience is insensitive to the inclusion of comprehensive vectors of parental socio-economic status and child endowments, including parental and child IQ and birth weight. This finding is consistent with the results in Chowdhury, Sutter and Zimmermann (2018), indicating that the transmission of time preferences is independent of socio-economic status.

As mentioned previously, only 2.8 percent of children were born when parents answered the time preference question. However, including those children in the analysis could be problematic, as having a child may affect revealed impatience. To test for this possibility, Column (6) replicates our preferred model while restricting the sample to those children born after the elicitation of parents' preferences. The results are robust to this restriction. Moreover, as discussed in Section 2, parents who planned having children in the near future might prefer the flat (impatient) wage profile simply to be able to afford having children and not because they truly were impatient. Therefore, Column (7) excludes parents (and their children) who had their first child before the patient wage profile would be fully implemented (i.e. before 1978). The results are again insensitive to this restriction, suggesting that our time preference measure is not just capturing correlations between parents' and children's fertility preferences. Lastly, Column (8) shows that the transmission is similar in the subsample of children with at least one sibling in the sample.

Table 2
Intergenerational Transmission of Impatience

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Impatient_p	0.063**	0.057*	0.073**	0.069**	0.069**	0.077**	0.076**	0.119***
	(0.030)	(0.030)	(0.031)	(0.031)	(0.031)	(0.031)	(0.034)	(0.036)
$\text{Impatient}_p \times$ Mother	0.030	0.020	-0.002	0.004	-0.013	-0.007	-0.024	-0.039
	(0.041)	(0.041)	(0.041)	(0.041)	(0.041)	(0.042)	(0.051)	(0.047)
Daughter		0.060**	0.057**	0.057**	0.068***	0.059**	0.070**	0.061**
		(0.024)	(0.025)	(0.025)	(0.025)	(0.025)	(0.027)	(0.030)
Sample	All	All	All	All	All	1974+	1978+	Siblings
Child demographics		✓	✓	✓	✓	✓	✓	✓
Parent demographics			✓	✓	✓	✓	✓	✓
Parent School FE			✓	✓	✓	✓	✓	✓
Grandparent SES				✓	✓	✓	✓	✓
Parent SES					✓			
Child endowments					✓			
Observations	3,101	3,101	3,101	3,101	3,101	3,014	2,197	2,255
Average	0.352	0.352	0.352	0.352	0.352	0.347	0.314	0.353

Standard errors in parentheses, clustered at the parent level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome variable indicates whether the child is impatient (1) or not (0). Each Column presents the results from separate regressions. *All* represents the full sample of children. The samples *1974+*, *1978+*, and *Siblings* only include respectively children born after 1973, children whose surveyed parent had their first child after 1977, and children with at least one sibling in the sample. Column (1) also control for parent gender. *Child demographics* include five-year age interval dummies, birth order dummies, and an indicator for being twin. *Parent demographics* include indicators for being born before 1954, born after 1954, female, child-parent gender combination, and birth order. *Parent School FE* include fixed effects for the surveyed parent's school in 1968. *Grandparent SES* includes grand-parental attitudes towards child education and work; an index for the grandparents' educational investment in the parent; quadratic taxable income in 1967 reported by the tax authorities; quadratic number of grandparents' children; indicators for the grandmaternal/paternal level of education, vocational training/education, grandmother/grandfather has subordinates, grandmother is housewife, gender of the surveyed grandparent, the parent lives with both parents at age 14, and indicators for missing observations for the different control variables. *Parent SES* includes the surveyed parent's spatial, verbal, and inductive abilities measured at age 14; the mother's and father's length of education, cumulated work experience through 2004, cumulated length of unemployment through 2004, the natural logarithm of average annual labor earnings 1980–2004; quadratic number of children; indicators for the child lives with both parents at age 16, the mother/father has children with another person than the parent, and missing observations for the different control variables. *Child endowments* include squared birth weight, child IQ, standardized length of highest completed education by 2016 by cohort; indicators for being born preterm and missing observations for the different control variables.

4.2 Gender Differences and the Role of Socio-Economic Status

Table 3
Childhood Socio-Economic Environment and the Intergenerational Transmission of Impatience by Child and Parent Gender

	Daughters			Sons		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: From Mothers to Children						
Impatient_p	0.168***		0.161***	-0.018		-0.024
	(0.046)		(0.048)	(0.047)		(0.048)
SES_p		-0.023	-0.004		-0.040**	-0.055**
		(0.021)	(0.025)		(0.020)	(0.023)
$\text{Impatient}_p \times \text{SES}_p$			-0.014			0.045
			(0.043)			(0.041)
Observations	843	843	843	757	757	757
Average of outcome	0.409	0.409	0.409	0.349	0.349	0.349
Panel B: From Fathers to Children						
Impatient_p	0.076		0.076	0.107**		0.102**
	(0.050)		(0.050)	(0.048)		(0.048)
SES_p		-0.033	-0.018		-0.023	0.004
		(0.030)	(0.034)		(0.028)	(0.034)
$\text{Impatient}_p \times \text{SES}_p$			-0.053			-0.060
			(0.061)			(0.055)
Observations	749	749	749	715	715	715
Average	0.352	0.352	0.352	0.288	0.288	0.288

Standard errors in parentheses, clustered at the parent level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome variable indicates whether the child is impatient (1) or not (0). Each Panel-Column presents the results from separate regressions. All models include *Child demographics*, *Parent demographics*, *Parent School FE*, and *Grandparent SES* (see the table note in Table 2 for details).

So far, we have documented a significant correlation of impatience across generations and found that daughters are overall more impatient than sons. Next, we study whether our full sample results are produced by certain child-parent gender pairs. To examine this question, we estimate three separate models for each parent-child gender pair presented in Table 3:[20] one depicting the impatience correlation

[20]The number of observations in this table does not sum to 3,101, because in few cases there are not more than one observation for each parent school-parent gender-child gender combination (which is necessary to include school fixed effects).

[Column (1) and (4)], one regressing child impatience on parental socio-economic status [Column (2) and (5)], and one including both parental impatience and socio-economic status [Column (3) and (6)]. All regressions control for the large set of covariates included in our preferred model in Table 2, Column (4).

The transmission of impatience from mothers to daughters is particularly strong [Panel A, Column (1)]: daughters of impatient mothers are 16.8 percentage points more likely to be impatient compared to daughters of patient mothers, representing an increase of 44 percent relative to the mean for daughters. Meanwhile, mothers' and sons' impatience are uncorrelated; the estimate is insignificant and small in magnitude [Panel A, Column (4)]. Although not statistically significantly different, the intergenerational transmission from fathers to children is smaller in magnitude than the one from mothers to children. The transmission from fathers to sons is more robust than the one from fathers to daughters. However, the magnitudes of the correlations between father-son and father-daughter pairs are not statistically significantly different from each other. These gender differences in the transmission are in line with the findings in Gauly (2016).

As discussed in Section 2, risk preferences might be associated with our measure of time preferences and thereby influence the estimated transmission of impatience. Therefore, Appendix Table A5 tests the robustness of the gender-specific impatience transmission by controlling for children's risk preferences. The results are similar to the ones in Table 3, suggesting that risk preferences are not an important confounding factor influencing our results on impatience propagation.

Considering a potential socio-economic gap in impatience, the socio-economic childhood family environment has overall only limited predictive power for explaining heterogeneity in children's impatience. Socio-economic status is only significantly correlated with sons' impatience in the sample of mothers and sons. However, the estimated *SES* gap is statistically significant when pooling the four samples due to the increase in statistical power (not reported). Thus, we fail to replicate a large gap in children's economic preferences by their socio-economic family background as the one found by Deckers et al. (2017). One reason might be that they study a smaller ($N = 435$), quite selected sample of relatively young children (age 7–9) in two specific German cities and that we consider a much more comprehensive non-self-reported array of socio-economic characteristics. Jointly including parental impatience and socio-economic status reveals that the intergenerational transmission of impatience is insensitive to the inclusion of our socio-economic index [Column (3) and (6)].[21] In other words, we do not find that the socio-economic family environment is an important mediator of the transmission as proposed by Deckers et al. (2017).

4.3 Comparing Siblings

Table 4
Heterogeneity in Intergenerational Transmission of Impatience:
Sibling Fixed Effects

	Mother			Father		
	(1)	(2)	(3)	(4)	(5)	(6)
I_p × Daughter	0.209**		0.256***	0.015		0.018
	(0.087)		(0.084)	(0.092)		(0.097)
I_p × Norm. Age		-0.011	0.003		0.019**	0.019
		(0.008)	(0.011)		(0.008)	(0.012)
I_p × Norm. Age			-0.027**			0.003
×Daughter			(0.014)			(0.017)
Daughter	0.006	0.064*	0.004	0.092**	0.099***	0.088**
	(0.044)	(0.038)	(0.043)	(0.040)	(0.036)	(0.041)
Observations	1,170	1,170	1,170	1,085	1,085	1,085
Average	0.390	0.390	0.390	0.314	0.314	0.314

Standard errors in parentheses, clustered at the parent level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome variable indicates whether the child is impatient (1) or not (0). Each Column presents the results from separate regressions. All models include sibling fixed effects and *Child demographics* (see the table note in Table 2 for details). *Norm. Age* refers to normalized age, which is calculated as the child's age in years subtracted 27, i.e. mean child age. All models except for Columns (1) and (4) further control for normalized age; Columns (3) and (6) also control for normalized age interacted with daughter. I_p denotes having an impatient parent.

In this last part of the analysis, we compare siblings within the same family to investigate whether parents differentially affect children depending on child gender and age. Specifically, can we say something more firmly about the gender-specific transmission of impatience found in the previous subsection and whether

[21]Given the finding in Dohmen et al. (2010) of a clear relationship between ability and impatience, we show in Appendix Table A6 that all results in Table 3 are similar when including the parent's IQ and *Child endowments*.

the transmission persists as children age? Table 4 reports these results for the samples of mothers and fathers, separately.[22]

Column (1) confirms our previous finding that mothers strongly influence daughters but not sons and can thus be interpreted as a causal differential transmission effect. Put differently, comparing sister-brother pairs shows that impatient mothers are considerably more likely to transmit their impatience to daughters relative to sons. Daughters of an impatient mother are 20.9 percentage points (55 percent relative to the mean of daughters) more likely to be impatient compared to sons of the same mother. Meanwhile, fathers do not differentially transmit impatience to sons relative to daughters [Column (4)], as foreshadowed in Table 3. The fact that we compare siblings within the same family allows us to exclude the childhood socio-economic environment and shared genetic factors as important mediators of the impatience transmission.[23]

These results indicate that role-modeling is an important mechanism for the transmission of impatience in mother-daughter dyads. In line with this, Brenøe and Lundberg (2017) show that, in terms of educational attainment, daughters benefit relatively more than sons from maternal education (and that sons benefit relatively more than daughters, but to a smaller extent, from paternal education). Similarly, Brenøe (2018) and Humlum, Nandrup and Smith (2017) show that mothers transmit gender norms to daughters but not to sons. This suggests that the differential transmission of maternal human capital to daughters relative to sons is due to differences in social interactions between mothers and their children and that especially mothers act as more prominent role models for daughters relative to sons.

To examine the permanence of the intergenerational transmission, we also explore whether the correlation differs by child age. Overall, for mothers, we do not observe any differential transmission by child age [Column (2)]. However, the results in Column (3) indicate that the differential effect from mothers to daughters relative to sons diminishes as daughters age. This is probably, in large part, because older children tend to be more impatient in our data [Appendix Figure A1].[24] For fathers [Columns (5) and (6)], the age pattern is different. Fathers affect older chil-

[22]All results in Table 4 are nearly identical when controlling for *Child endowments* (not reported).
[23]The results are similar when restricting the sample to full siblings.
[24]This might be due to how the question was posed. Note that *Norm. Age* refers to normalized age, which is calculated as the child's age in years subtracted 27, i.e. mean child age.

dren more strongly relative to younger siblings. Although the estimate in Column (6) is not statistically significant, it is similar in magnitude to the one in Column (5). Consequently, these findings emphasize the persistence of the transmission effect, stressing that it does not fade out as children age.

5 Conclusion

We document that parents transmit their time preferences to children, that the transmission is particularly strong from mothers to daughters, and that the intergenerational propagation of impatience does not fade out as children age. Our setup for this study is unique for three reasons, allowing for important contributions to the literature on intergenerational transmission of preferences. First, parents' and children's impatience are measured four decades apart, eliminating any concerns regarding reverse causality. Second, we link the survey data to high-quality administrative data, enabling a rigorous examination of the interplay between socio-economic status and preference transmission. From this analysis, we conclude that, in our setting, the childhood socio-economic family environment does not mediate the propagation of impatience. Third, for the majority of parents, we observe several children, making a mechanism analysis possible in which we compare siblings within the same sibship. Our findings favor the explanation that the social environment (nurturing) affects children's impatience and, in particular, role-modeling between mothers and daughters. Yet, future research needs to explore what exact factors in parental care are responsible for this transmission.

Our results have implications for both policy making and economic modeling. First, the fact that the transmission is particularly strong and long-lasting from mothers to daughters may hint at a possible channel for implementing policies. The existing literature and the correlations between our measure and real-world economic outcomes stress that time preferences indeed predict important lifetime outcomes. Therefore, if the intergenerational transmission of impatience translates into a higher likelihood of ending up in debts, earning lower labor incomes, and having worse health for the children of impatient parents, then policy makers may have a clear incentive to intervene. To break up the positive correlation between parents' and children's impatience, effective policies would target young children's education and nurturing. Second, macroeconomic models considering multiple generations usually assume that time preferences propagate from parents to offspring (see e.g. Krusell and Smith (1998)). The empirical evidence on the intergenerational transmission of preferences, however, has previously only considered relatively short time horizons. Our study provides support for the assumption in macroeconomic models that time preferences indeed transmit from generation to generation and that this propagation persists over a very long time period.

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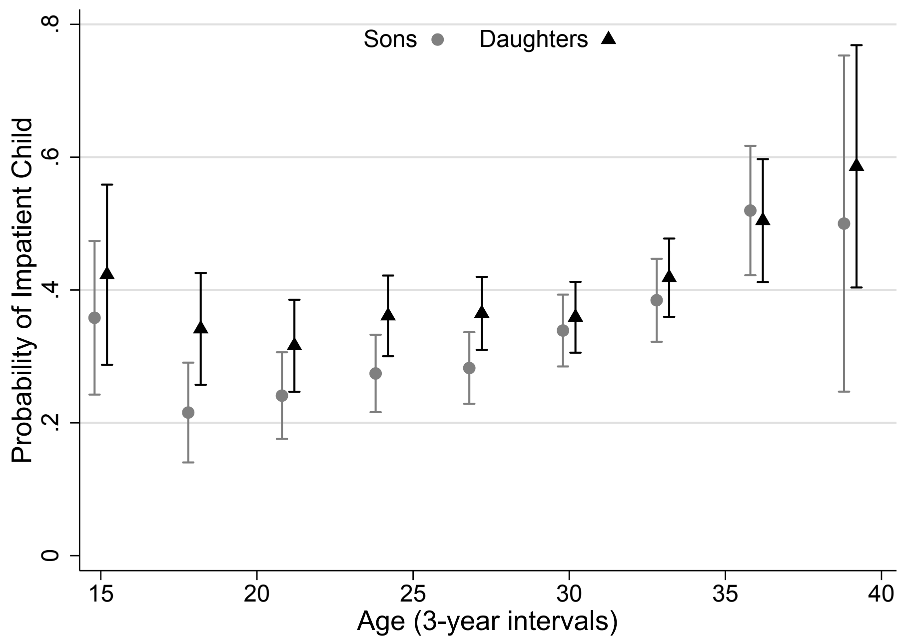
A Appendix

Figure A1
Child Impatience by Age and Gender



Note: This graph illustrates the share of impatient children by age and gender. Age is shown in 3-year intervals. The whiskers represent the 95 percent confidence interval.

Table A1
Literature on Transmission of Time Preferences

Article	Type of Measure	Measure	Sample (#, age)	Delay between parent and child measurement	Main findings	Comment
Bartling et al. (2010)	Experimental	Intertemporal choices over money (mothers; delays of 6 and 12 months) and gummy bears (children; later today, tomorrow, or the day after)	270 children (age 5–6) and their mothers, i.e. no fathers	Simultaneous measurement in separate rooms	Children of more patient mothers are more likely to be patient. Only significant correlations for the near-present tradeoffs (see Kosse and Pfeiffer (2013)).	Only weak evidence; small sample
Brown and van der Pol (2015)	Survey question	Question on planning horizon as a proxy for time preferences	Panel data from Household Income Labour Dynamics of Australia (HILDA), 6 waves; children: 2757 (male) + 2555 (female); parents: 2965 mothers + 2338 fathers; analysis restricted to young adults (age 16–25) and both parents; examine all four dyads	Have data from 6 waves over 8 years; compare transitions in answer categories from one to next year and find relatively stable responses; do not explore persistence of transmission, however.	Support for transmission of time preferences; gender differences: association of time preferences larger for mothers than fathers	Hypothesize correlation of planning horizon and discount rate
Chowdhury, Sutter and Zimmermann (2018)	Experimental	Choice lists with tradeoffs next day vs 3 weeks (children), 3 months (all) or 1 year (parents)	Household sample from Bangladesh; relatively poor families; 911 children (age 6–17); 544 pairs of mothers/fathers	Simultaneous measurement in separate rooms	Significant correlation between mothers', fathers', and children's preferences; correlation of similar size for both genders	Relatively homogenous sample; SES has only limited predictive power for children's preferences
Gauly (2016)	Survey question	Patience question of the German Socio-Economic Panel (SOEP)	2395 "children" for whom it was possible to identify biological parents; age not reported	Simultaneous measurement (same year of SOEP)	Parents transmit own attitudes to children via direct socialization. Find lowest correlation (of all measures) for patience, but large correlations between father-son and mother-daughter pairs.	Included also a measure of reciprocity and examine the persistence of the correlation across five years. Find weaker correlations when delay increases.
Kosse and Pfeiffer (2013)	Experimental	See comment	See comment	See comment	Mothers' and children's preferences for immediate gratification (present bias) are positively correlated. No significant correlation between mothers' and children's impatience.	Use data described in Bartling et al. (2010).

This table restricts attention to studies eliciting time preferences or proxies of these. There is a larger literature focusing on other preference domains (see Section 1.).

Table A2

Associations Between Impatience and Socio-Economic Status

	(1)	(2)	(3)	(4)	(5)	(6)
	Parents by age 50					Child
	SES Index	Education (years)	Unemployment	Work Experience	Log(Earnings)	Education (years)
			Women			
Impatient	-0.17**	-0.46**	0.82***	-2.58***	-0.44***	-0.50***
	(0.07)	(0.20)	(0.17)	(0.56)	(0.13)	(0.11)
Observations	1,369	1,369	1,369	1,369	1,369	1,608
Average	-0.005	12.181	1.863	21.012	11.450	14.176
			Men			
Impatient	-0.22***	-0.60***	0.42***	-0.75	-0.26**	-0.52***
	(0.06)	(0.17)	(0.15)	(0.57)	(0.11)	(0.13)
Observations	1,370	1,370	1,370	1,370	1,370	1,490
Average	0.005	12.943	1.322	21.814	11.916	13.634

Standard errors in parentheses, clustered at the parent school level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Each Column-Gender presents the results from separate regressions. The sample of parents correspond to the original DLSY respondents who have at least one child. The sample of children are children to the DLSY parents. All models include *Parent demographics*, *Parent School FE*, and *Grandparent SES* (see the table note in Table 2 for details). Column (6) also include *Child demographics*. *SES Index* (standardized with mean zero and standard deviation one) is the first principal component from a principal component analysis; see Appendix Table A3. *Education* measures the length of highest completed education in years by 2016. *Unemployment* measures the cumulated length of unemployment in years 1980–2004. *Work Experience* measures the cumulated length of work experience in years 1964–2004. *Log(Labor Earnings)* is the natural logarithm of average annual labor earnings 1980–2004.

Table A3
Principal Component Analysis: SES Index

	First Component	Average
Mother's Education (years)	0.27	12.63
Father's Education (years)	0.32	12.97
Mother's Education missing	-0.21	0.0017
Father's Education missing	-0.26	0.0045
Parent's # of Children	-0.12	2.47
Mother has children with other than Father	-0.21	0.14
Father has children with other than Mother	-0.14	0.17
Mother's years of unemployment 1980–2004	-0.30	2.01
Mother's years of work experience 1964–2004	0.37	20.58
Father's years of unemployment 1980–2004	-0.22	1.17
Father's years of work experience 1964–2004	0.25	23.36
Log(Mother's mean labor earnings 1980–2004)	0.40	11.69
Log(Father's mean labor earnings 1980–2004)	0.38	12.18
	Eigenvalue	**Proportion**
Component 1	2.87	0.22
Component 2	1.59	0.12
Component 3	1.41	0.11
Component 4	1.25	0.10
Component 5	1.05	0.08
Component 6	0.92	0.07
Observations	3,518	

Principal component analysis (PCA) of the socio-economic status experienced during children's childhood. The sample includes all children born by 1996, also those who did not answer the survey in 2010. We use the first component to construct the SES index.

Table A4

Associations Between Impatience and Fertility Preferences

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Fertility Preferences at Age 22 and Early Fertility						
	Desired # of Children			Has Any Children by		
	0	1	2	1973	1976	1979
			Women			
Impatient	-0.01	0.02	-0.00	0.06***	0.15***	0.10***
	(0.02)	(0.02)	(0.03)	(0.02)	(0.03)	(0.03)
Observations	1,267	1,267	1,267	1,369	1,369	1,369
Average	0.093	0.066	0.539	0.071	0.264	0.496
			Men			
Impatient	0.07***	-0.01	-0.07**	0.00	0.02	0.04
	(0.02)	(0.02)	(0.03)	(0.01)	(0.02)	(0.03)
Observations	1,211	1,211	1,211	1,370	1,370	1,370
Average	0.116	0.043	0.597	0.012	0.087	0.231
Panel B: Complete Fertility by 2016 (Age 62)						
	Has Any Child	# of Children	Age at First Birth	Age at Last Birth	# of Child w *I* Obs	Daughter w *I* Obs
			Women			
Impatient	-0.01	0.04	-1.50***	-0.95**	0.09	0.03
	(0.02)	(0.07)	(0.35)	(0.39)	(0.07)	(0.04)
Observations	1,369	1,369	1,191	1,191	1,191	963
Average	0.870	1.836	25.653	28.919	1.358	0.669
			Men			
Impatient	-0.09***	-0.21***	-0.77*	-0.48	-0.00	-0.01
	(0.03)	(0.08)	(0.43)	(0.44)	(0.07)	(0.04)
Observations	1,370	1,370	1,061	1,061	1,061	871
Average	0.778	1.682	28.900	31.884	1.396	0.659

Standard errors in parentheses, clustered at the parent school level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Each Panel-Column-Gender presents the results from separate regressions. The sample includes all original DLSY respondents. All models include *Parent demographics*, *Parent School FE*, and *Grandparent SES* (see the table note in Table 2 for details). *Desired # of Children* indicates whether the respondent in 1976 reported that their desired number of children was respectively, 0, 1, or 2, leaving 3 or more children the omitted category (due to the survey question, those who already had children and did not want more have a coded desired number of children equivalent to the number of children they had by 1976; the correlations between impatience and desired fertility is similar when excluding those who already had children in 1976). *Has Any Children by* indicate whether the respondent had at least one child by 1973, 1976, and 1979, respectively. *Has Any Child* indicates whether the person has any children by 2016. *# of Child w I Obs* measures the parent's number of children with an observation on impatience, conditional on having at least one child by 1996. *Daughter w I Obs* indicates whether the parent has at least one daughter with an observation on impatience, conditional on having at least one child in the survey.

Table A5
Children's Time and Risk Preferences: The Dependent Variable is Child Impatience

	Daughters				Sons			
	%	(1)	(2)	(3)	%	(4)	(5)	(6)
Impatient_p			0.178***	0.070			-0.025	0.099**
			(0.046)	(0.052)			(0.048)	(0.049)
Lottery Ticket Price 0–20,000 DKK								
50–499	52	0.001	-0.021	-0.003	41	-0.068*	-0.050	-0.050
		(0.030)	(0.044)	(0.047)		(0.037)	(0.054)	(0.061)
500–1999	11	-0.060	-0.127**	0.015	27	-0.111***	-0.119**	-0.061
		(0.045)	(0.064)	(0.073)		(0.039)	(0.060)	(0.065)
2,000-20,000	5	-0.066	-0.101	-0.055	13	-0.108**	-0.163**	-0.024
		(0.060)	(0.086)	(0.116)		(0.048)	(0.071)	(0.078)
Investment 0–500,000 DKK (shown in 1,000 DKK)								
0.001–1	14	0.048	0.056	0.032	12	-0.002	0.065	-0.038
		(0.041)	(0.060)	(0.065)		(0.044)	(0.067)	(0.063)
1–249	30	-0.058*	-0.054	-0.066	26	-0.111***	-0.081	-0.117**
		(0.031)	(0.050)	(0.049)		(0.032)	(0.052)	(0.048)
250–500	12	-0.119***	-0.133**	-0.180***	16	-0.079**	-0.037	-0.096*
		(0.042)	(0.063)	(0.066)		(0.038)	(0.060)	(0.056)
General Risk 1–10 (1 = avoid risks; 10 = do not mind taking risks)								
4–7	63	0.028	-0.009	0.127**	61	-0.088**	-0.033	-0.135**
		(0.037)	(0.054)	(0.056)		(0.040)	(0.056)	(0.066)
8–10	20	0.065	0.023	0.161**	24	-0.114**	-0.008	-0.194***
		(0.045)	(0.066)	(0.069)		(0.045)	(0.064)	(0.071)
Sample		All	Mother	Father		All	Mother	Father
Observations		1,571	827	744		1,456	745	711
Average		0.382	0.409	0.352		0.317	0.346	0.286

Standard errors in parentheses, clustered at the parent level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome variable indicates whether the child is impatient (1) or not (0). Each Column presents the results from separate regressions. All models include *Child demographics*, *Parent demographics*, *Parent School FE*, and *Grandparent SES*, (see the table note in Table 2 for details). The *Lottery Ticket Price* question is: *You have the opportunity to buy a lottery ticket. There are 10 people in the lottery. The prize is 20,000 DKK. The winner of the lottery is found by lottery, i.e. everyone has the same chance of winning. What price do you want to pay for a lottery ticket for this lottery?* The *Investment* question is: *You have won 500,000 DKK in the lottery! You are contacted by a reputable bank that offers you an investment opportunity. The terms are as follow: You have a 50 percent probability of doubling your investment within two years. However, there is also a 50 percent probability of losing your investment. How much of the 500,000 DKK will you invest?* The *General risk* question is: *Do you perceive yourself as a person willing to take risks to achieve something in life, or avoid any risks? Answer on a scale from 1–10, where "1" means avoiding risks and "10" means you do not mind taking risks.* % indicates the percentage of respectively daughters and sons answering within the range indicated for each of the three questions.

Table A6
Childhood Socio-Economic Environment and the Intergenerational Transmission of Impatience by Child and Parent Gender: Controlling for Parental IQ and Child Endowments

	Daughters			Sons		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: From Mothers to Children						
Impatient$_p$	0.153***		0.151***	-0.055		-0.055
	(0.046)		(0.047)	(0.048)		(0.048)
SES$_p$		-0.007	0.006		-0.029	-0.045**
		(0.021)	(0.021)		(0.019)	(0.022)
Impatient$_p$ $\times$ SES$_p$			-0.013			0.045
			(0.044)			(0.037)
Observations	843	843	843	757	757	757
Average of outcome	0.409	0.409	0.409	0.349	0.349	0.349
Panel B: From Fathers to Children						
Impatient$_p$	0.071		0.071	0.107**		0.105**
	(0.050)		(0.050)	(0.048)		(0.048)
SES$_p$		-0.008	0.002		-0.007	0.025
		(0.030)	(0.033)		(0.032)	(0.035)
Impatient$_p$ $\times$ SES$_p$			-0.040			-0.076
			(0.058)			(0.060)
Observations	749	749	749	715	715	715
Average	0.352	0.352	0.352	0.288	0.288	0.288

Standard errors in parentheses, clustered at the parent level. * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The outcome variable indicates whether the child is impatient (1) or not (0). Each Panel-Column presents the results from separate regressions. All models include *Child demographics*, *Parent demographics*, *Parent School FE*, *Grandparent SES*, *Child Endowments*, the surveyed parent's spatial, verbal, and inductive abilities, and an indicator for missing parental ability (see the table note in Table 2 for details).