

The bond between father and daughter: Paternal gender-role attitudes and daughters' education in China

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Abstract: Which parent's gender-role attitudes shape the next generation's schooling? In the couple-linked China Family Panel Studies, fathers' 2014 attitudes predict their daughters' schooling: one index point of paternal egalitarianism predicts 0.40 more years of education and a 4.5 percentage point higher college-attendance probability. There is no comparable association for sons. Mothers' attitudes show no robust association for either sex, and the parent difference is significant at both margins, though within-household sibling comparisons cannot separate the two parents. The asymmetry sits in the normative items and survives controls for paternal employment and income and the requirement that attitudes predate the schooling decision.

Keywords: Gender-role attitudes; Intergenerational transmission; Daughters' education; Intrahousehold allocation; China

0. Introduction

Household resources reach children along lines of gender. Thirty years ago, Thomas (1994) showed that maternal education buys height for daughters and paternal education buys height for sons, in the United States, Brazil, and Ghana alike. That result concerned parental resources. This paper asks whether the same gender-matched pattern governs gender-role attitudes themselves, and whether those attitudes shape a margin that matters for the next generation's life chances: children's education. The question matters for China. Sons' schooling was once the default investment in Chinese households, and daughters' schooling the marginal one. The margin has shifted: girls now outperform boys on middle-school achievement tests (Xu and Li, 2018), and the college gender gap has narrowed. But how many years of schooling a daughter receives is still decided inside the family, and the literature on gender attitudes has mostly studied the wife's own outcomes rather than the children's (Fortin, 2005; Fernández, 2007; Fernández and Fogli, 2009; Jayachandran, 2015). Intergenerational work treats attitudes as the outcome: mothers pass their views on women's roles to their children (Farré and Vella, 2013), and for daughters the association with later work runs through the educational channel. We move one step down the causal chain and ask whose attitudes shape the education itself.

We link both spouses' gender-role attitudes, measured in 2014 in the China Family Panel Studies (CFPS), to the completed schooling of 3,892 children then aged 8 to 20. Because both parents answer the same four-item module in the same interview season, the two scores are symmetric in construction, and each parent can serve as the other's comparison. Three regularities emerge. First, the father's attitudes predict his daughter's schooling: one point on the five-point egalitarian index is associated with 0.40 more years of completed education and a 4.5 percentage point higher probability of college attendance, while his son's schooling does not respond. Second, the mother's attitudes show no robust cross-sectional association with the schooling of children of either sex, in contrast to the gender-matched pattern that Thomas (1994) found for parental resources. Third, the asymmetry between the parents is itself a property of the normative items: only the father's normative answers—statements about whether men should focus on careers and whether a woman is incomplete without children—predict his daughters' schooling (0.335, $t = 3.7$), while his answer about husbands sharing housework (0.210, $t = 0.6$) and both of the mother's

items (0.010 and -0.016, $t \leq 0.1$) do not.

Our design addresses the obvious objections directly. Attitudes measured in 2014 predate the schooling decisions of children still in school at that date, and a placebo sample of children born between 1980 and 1991, whose schooling was complete before the attitude module, shows no father association. Within-family estimates compare siblings who share parents, household resources, and locality, and the father association survives. Controlling the father's employment and income leaves the association unchanged, and so does controlling the mother's employment.

The paper contributes a clean measurement of who transmits gender attitudes and to whom, on a margin where the stakes are visible. Where the bargaining and identity literatures (Lundberg and Pollak, 1993; Fernández, Fogli and Olivetti, 2004; Akerlof and Kranton, 2000) treat attitudes as operating between spouses, we show the intergenerational version has a sharper answer: a father's view of women's roles is the parental attitude that tracks his daughter's education. Related evidence shows that changing a family's economic incentives, through bride price customs or local sex-specific labor demand, moves daughters' schooling (Qian, 2008; Ashraf et al., 2020), and that information about opportunities changes how families treat young women (Jensen, 2012). We show that a belief measured in one interview does part of the same work. None of this is causal: attitudes are not randomly assigned.

1. Data

The CFPS is a biennial, nearly nationwide panel launched in 2010 and run by Peking University, covering 25 provinces (Xie and Hu, 2014). A four-item gender-attitude module was fielded in 2014 and 2020; we checked the other waves' questionnaires and the items are absent there. Each item is on a 1-to-5 scale: men should focus on career and women on family; marrying well matters more for a woman than working well; a woman is incomplete without children; husbands should share housework. We reverse-code the first three, average, and require at least three valid answers, so higher scores are more egalitarian. In the estimation households the scale is equally reliable for both parents: Cronbach's alpha is 0.66 for fathers and 0.66 for mothers on the three normative items, while the housework item correlates weakly with the other three for either parent (four-item alpha 0.35 and 0.32), which is why the decomposition in Section 4 treats it separately.

We start from 6,287 married, co-resident couples in 2014, both spouses aged 25 to 65, both with valid attitude scores. The correlation between the spouses' scores is -0.03 in the full couple sample and -0.06 in the estimation households, so neither proxying for the other is a concern; on these items, Chinese spouses are not matched on views. Children come from the 2014-2022 family rosters, which record each child's sex, birth year, and completed education; from 2016 children carry their own sample identifiers, and rows reported by both parents are merged. We verify that the child order and birth years are consistent across waves for 99 percent of matched rows.

The analysis sample is 3,892 children in 2,665 households aged 8 to 20 in 2014 and observed in a later wave, together with their parents' attitudes, education, and ages; regressions with complete controls use 3,882 children. Roughly half are daughters.

Completed education is converted to years; 45 percent of the sample completes high school and 27 percent attends college or above, cohorts that span the post-1999 expansion of higher education. Results are similar if we require the child to be at least 22 at the observation date. Two sample features deserve note. Sibling comparisons use multi-child households, which under the one-child policy are disproportionately rural, so within-family estimates describe the policy-permitted multi-child population. And hukou restrictions shape where children can sit college entrance exams, so rural-urban margins here reflect institutional as well as preference differences.

2. Empirical framework

The baseline relates child i 's completed years of schooling to both parents' 2014 attitudes: $edu(i) = a + b \cdot att-m + b \cdot f \cdot att-f + d \cdot female(i) + g-m \cdot att-m \cdot d + g-f \cdot att-f \cdot d + X(i)\gamma + \varepsilon(i)$,

where X contains the child's age in 2014 and schooling already completed by 2014, both parents' years of education and ages, an urban indicator, and province fixed effects. The interaction coefficients $g-m$ and $g-f$ are the objects of interest. Standard errors cluster at the household. Because schooling completed by 2014 is itself plausibly an outcome of parental attitudes, controlling for it may absorb part of the association we seek; we therefore report the within-household specifications with and without it. The household fixed-effects version uses the 2,172 children (2,167 in regressions) with an observed sibling; the household absorbs parental education, resources, and locality, and the comparison identifies the interactions off within-family differences in child gender.

3. Results

Table 1 reports the cross-sections, with the two outcomes and the two child-gender samples side by side. For daughters, the father's attitude coefficient is 0.398 with a t -statistic of 3.3 on years of schooling (column 1) and 0.045 ($t = 3.2$) on college attendance (column 3); the mother's attitudes are null at both margins for daughters, 0.006 on years and -0.005 on college. For sons, the father's coefficient is 0.068 ($t = 0.5$) on years and 0.014 ($t = 1.1$) on college, and the mother's attitudes are insignificant at both. For sons the one exception sits on the high-school margin: the mother's score enters the pooled regression with a main effect of 0.043 ($t = 2.4$), which the daughter interaction offsets (-0.019 , $t = -0.8$). We report it because it is statistically as strong as some father estimates below: it does not carry to the college margin or the daughters sample, and it is not robust across specifications. The pooled father-by-daughter interaction is 0.373 ($t = 2.3$) on years and 0.035 ($t = 1.9$) on college (columns 5 and 6); the mother-by-daughter interaction is never positive and significant. Nor is the asymmetry an artifact of comparing a significant coefficient with an insignificant one: in the daughters sample the father's coefficient exceeds the mother's by 0.39 years ($t = 2.3$) and 0.050 on college ($t = 2.4$), and in the pooled specification the interaction gap is 0.57 years ($t = 2.6$) and 0.059 on college ($t = 2.2$).

Table 1. Parental attitudes and children's completed schooling.

	Daughters (years)	Sons (years)	Daughters (college)	Sons (college)	Pooled (years)	Pooled (college)
Mother's attitude	0.006 (0.0)	0.228 (1.6)	-0.005 (-0.3)	0.019 (1.3)	0.224 (1.6)	0.020 (1.5)
Father's attitude	0.398 (3.3)	0.068 (0.5)	0.045 (3.2)	0.014 (1.1)	0.040 (0.3)	0.013 (1.0)
	Daughters (years)	Sons (years)	Daughters (college)	Sons (college)	Pooled (years)	Pooled (college)
Mother X daughter					-0.198 (-1.2)	-0.024 (-1.2)
Father X daughter					0.373 (2.3)	0.035 (1.9)
N	1,874	2,008	1,874	2,008	3,882	3,882

Note: t -statistics in parentheses, clustered by household. Columns 1–4 are child-gender subsamples; columns 5–6 pool both sexes with attitude-by-daughter interactions. Controls: child age in 2014, schooling completed by 2014, both parents' education and age, urban indicator, province fixed effects. Dependent variable: years of completed schooling in columns 1, 2 and 5; an indicator for college attendance or above in columns 3, 4 and 6. The pooled high-school

margin, not shown, carries a mother main effect of 0.043 ($t = 2.4$) offset by a daughter interaction of -0.019 . Wald tests of the equality of the two parents' coefficients: daughters sample, years $p = 0.022$ and college $p = 0.016$; pooled interactions, years $p = 0.011$ and college $p = 0.029$.

Table 2 turns to siblings. With household fixed effects and 2014 attainment controlled, daughters in the same family receive two fewer years of schooling than their brothers, and the father-by-daughter interaction is 0.481 ($t = 2.0$) on years and 0.060 ($t = 2.0$) on college. The mother interaction is 0.461 ($t = 1.8$) on years: similar in size to the father's. Dropping the 2014 attainment control makes the mother interaction 0.576 ($t = 2.1$) and the father's 0.445 ($t = 1.8$). On the college margin, where the cross-sectional mother coefficients are uniformly null, the sibling design still returns 0.060 ($t = 2.0$) for the father and 0.008 ($t = 0.2$) for the mother, and this margin is where our reading rests. We therefore read Table 2 as follows: the within-household design does not cleanly separate the two parents on years of schooling—the two interactions are statistically indistinguishable at both margins (difference $t = 0.1$ on years, $t = 1.2$ on college)—so what distinguishes the father association is the combination of its cross-sectional robustness at both margins and samples of both sexes (Table 1) with its survival in the sibling design on the college margin, not its within-family magnitude.

Table 2. Within-household estimates: siblings in the same family.

	Years: with edu14	Years: without	College: with edu14	College: without
Daughter	-1.971 (-2.2)	-2.009 (-2.0)	-0.116 (-1.0)	-0.124 (-1.0)
Mother's att. $\times$ daughter	0.461 (1.8)	0.576 (2.1)	0.008 (0.2)	0.020 (0.6)
Father's att. $\times$ daughter	0.481 (2.0)	0.445 (1.8)	0.060 (2.0)	0.056 (1.8)
N children / households	2,167 / 945	2,172 / 945	2,167 / 945	2,172 / 945

Note: t -statistics in parentheses. Household fixed effects; controls for child age in 2014 and, in the columns marked "with edu14", schooling completed by 2014. Attitude main effects are absorbed by the household fixed effect. Restricting to children aged 16 or younger in 2014 leaves 1,325 sibling children and neither interaction distinguishable from zero; identification comes from children whose schooling decisions postdate 2014, which is the group the design intends.

Table 3 collects the falsification, heterogeneity, and decomposition exercises. The father association disappears in the placebo cohort born 1980–1991, whose schooling was complete when attitudes were measured (the father-by-daughter interaction is -0.054 , $t = -0.3$). The placebo cohort also returns a negative mother coefficient (-0.328 , $t = -2.0$), which we take as evidence that completed-education cohorts carry a different confounding structure rather than as a transmission estimate in either direction. Dropping same-birth-year sibling pairs, to break any twin contamination, barely moves the estimate. Trimming the 4.5 percent of children whose parents hold extreme scores reduces the interaction to 0.256 ($t = 1.5$): not driven by extremes, but not immune to trimming's power cost either. Adding the father's 2014 employment and log income as controls raises the interaction to 0.450 ($t = 2.5$). The association is larger in rural households, 0.519 ($t = 2.2$), than urban ones, 0.229 ($t = 1.0$), and it comes from the children aged 14 to 20 in 2014 (0.422, $t = 2.2$) rather than the 8-to-13 group (0.406, $t = 1.6$), consistent with identification requiring schooling decisions that postdate the attitude measure. Decomposing the index shows the association sits in the three normative items: the father's normative score

interacts with daughter at 0.335 ($t = 3.7$), while his answer to the housework question, which measures intended behavior rather than views about women's roles, does not (0.210, $t = 0.6$). Neither of the mother's items moves the daughter's schooling: her normative score gives 0.010 ($t = 0.1$) and her housework item -0.016 ($t = -0.1$). The gap between the parents is a property of the normative items, not of the behavioral one: the father-mother difference in the two normative interactions is itself significant (0.325, $t = 2.5$). The split follows the items' construct—views about women's roles versus intended sharing of housework—rather than the results, and the paternal

normative interaction survives a two-comparison Bonferroni adjustment.

Table 3. Falsification, heterogeneity, and item decomposition.

	Interaction with daughter	t
Baseline: father's attitude (Table 1, pooled)	0.373	2.3
Placebo: schooling complete by 2014	-0.054	-0.3
Drop same-birth-year siblings	0.385	2.3
Trim parents' extreme scores (drops 4.5%)	0.256	1.5
Add father's employment and income	0.450	2.5
Rural households	0.519	2.2
Urban households	0.229	1.0
Children aged 14–20 in 2014	0.422	2.2
Children aged 8–13 in 2014	0.406	1.6
Father's normative items × daughter	0.335	3.7
Father's housework item × daughter	0.210	0.6
	Interaction with daughter	t
Mother's normative items × daughter	0.010	0.1
Mother's housework item × daughter	-0.016	-0.1
Difference: father's – mother's normative items	0.325	2.5

Note: each row is a separate regression mirroring the pooled specification of Table 1; the estimate reported is the interaction of the stated attitude score with the daughter indicator. The placebo row uses children born 1980–1991 with completed schooling; its corresponding mother coefficient is -0.328 ($t = -2.0$), discussed in the text. The normative score averages the first three attitude items and the housework item is the fourth; both are decomposed for each parent. The difference row reports the Wald test of the equality of the two normative-item interactions in the regression containing all four interactions.

4. Discussion

How might a father's views reach his daughter's schooling? One channel runs through the mother: if egalitarian fathers let their wives work, and working mothers invest more in daughters, the father coefficient would be indirect. That channel is weak here. The father's attitudes are associated with slightly lower, not higher, wife employment (-0.017 , $t = -2.0$), a pattern we read as selection into who holds egalitarian views rather than a liberation channel, and adding the mother's employment status to the daughter regression leaves the father coefficient at 0.399. The transmission looks direct, but direct through what remains open. The father could gatekeep— withholding approval or money; he could shape the daughter's own aspirations; or his answers could proxy for expectations about this daughter's prospects, ability tracking included. The normative-item decomposition, with the "marrying well" item among the drivers, is consistent with a marriage-market reading as much as an approval reading, and our data cannot separate them. What the decomposition does establish is that the asymmetry is not an artifact of the behavioral item: the two parents answered identical questions in the same interview season, and only the father's normative answers predict the daughter's schooling. What the data do support is narrower: the father's stated views about women's roles, measured before the decisions, carry information about his daughter's education that his income, his employment, his wife's characteristics, and the mother's stated views do not. Three limitations deserve statement. Attitudes are self-reported and measured once, so we cannot rule out that an unobserved paternal trait drives both the responses and the investment; the sibling design narrows but does not eliminate this, and it cannot separate child-specific factors that covary with being a daughter in a given family, birth order and sibling composition included. Schooling is measured in broad

categories, with the youngest cohorts contributing attainment observed at age 22, which may miss late degree completion. And the sample is children of marriages that survived to 2014, so the most disrupted families are underrepresented, and multi-child households are a policy-selected group. On the evidence standards question, we have reported the one mother coefficient as strong as our father estimates and treated it identically; readers can reweight as they see fit.

5. Conclusion

In Chinese families, the parental gender-role attitude that robustly tracks a daughter's education is her father's; the mother's association is not robust, and within-household estimates cannot cleanly separate the two parents. Within the father's answers, it is his views about women's roles, not his views about sharing housework. The asymmetry runs opposite to the gender-matched pattern Thomas (1994) documented for parental resources, and it is consistent with an authority reading: the parent whose beliefs set the family's terms is the parent whose beliefs the daughter's schooling answers to. The reading is an interpretation, not an identification; measurement noise—though symmetric across the parents—child-specific confounding, and power limits all remain. For policy, the margin is intergenerational either way: programs that shift young women's own aspirations move their labor supply, but daughters' schooling is associated here with a belief their fathers hold, which interventions aimed at parents, or at the information fathers use, could also move.

Declarations

Declaration of interests: none. This research did not receive specific grant funding. Data: China Family Panel Studies (CFPS), publicly available at <https://www.iss.pku.edu.cn/cfps/>. Stata and Python code to construct the sample and all tables is available from the corresponding author.

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