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Intergenerational transmission of education: The case of rural China

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A B S T R A C T

The intergenerational transmission of education has received considerable attention in recent empirical research in many countries. However, the research on intergenerational transmission of education in China is still relatively rare. This paper investigates the impact of parental schooling on their children's schooling in rural China using the data collected by the authors themselves. Our results show that (i) the intergenerational transmission of education in rural China is not as high as those have been reported in the literature for several other countries;(ii) There exists significant transmission effect of education in the subgroup born after the 1980s, but not for those who were born in the year of 1980 onward. The results also stand up to several different tests and robustness checks. Our findings suggest that promoting the equal education opportunity and investing in children of disadvantaged group will have long-term effects for the accumulation of human capital. China can promote increasing gains for its acquisition of human capital, and tap into this foundation for sustainable growth and development in the future.

1. Introduction

Intergenerational transmission has long been a central topic in social sciences. This concerns what is transferred down from one generation to its descendants. Many studies show that a considerable amount of socioeconomic inequality originates from disparate transmission effects across demographics (Becker, 1994; Li, 2006). There has been a substantial quantity of literature concentrated on the intergenerational transmission of income, which focuses on the correlation between the incomes of parents and their children (Björklund & Jäntti, 2009; Blanden, 2013; Corak, 2006; Deng, Gustafsson, & Li, 2013). Almost all of these studies show that there exists obvious intergenerational persistence of income.

The effect of education on inequality and intergenerational mobility of income has been analyzed in some literature. For example, some scholars found that parental education plays one of the most important roles in conferring economic status (Gong, Leigh, & Meng, 2012; Li & Zhou, 2014). More recently, Yang and Qiu (2016) argued that family investment in early education plays an important role in explaining income inequality and intergenerational income mobility. Furthermore, parents' educational attainment provides an important source of predictors, both short- and long-term, for the lives of their children (Black & Devereux, 2010).

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Increasing numbers of papers have emerged during the last ten years which focus on the intergenerational transmission of education (Hertz et al., 2007; Holmlund, Lindahl, & Plug, 2011; Maurin & McNally, 2008). However, compared to the abundant amount of studies on intergenerational transmission of income (Black & Devereux, 2010; Qin, Wang, & Zhuang, 2016), intergenerational transmission of education is a topic that has been less explored. This is especially true in the context of China. These heretofore neglected topics provide the impetus for our own research.

The overall goal of this paper is to examine the intergenerational transmission of education in the context of rural China. Under this goal, we have two specific objectives. First, we seek to investigate whether there exists any intergenerational transmission of education by using a unique dataset from rural China collected by the authors themselves. Second, we want to understand whether there exist any differences in intergenerational transmission of education across different birth cohorts.

The results based on ordinary least squares (OLS) show there exists significant transmission of education across generations. However, the intergenerational transmission effects estimated by family fixed effects (FFE) are comparatively minute. The results also show a significant transmission of education across generations for those cohorts born after 1980 but not for those cohorts born by 1980. This means the initial results were mainly driven by the cohorts born after 1980.

The rest of the paper is structured as follows. We present a literature review in Section 2. Section 3 introduces the data and identification strategy used in this paper. Section 4 presents the empirical results. The summary of our findings and some discussions are presented in the final section.

2. Previous studies

There are a number of studies which concentrate on the intergenerational relationship of education in many countries using different identification strategies and data sources. For example, Haveman and Wolfe (1995) compiled a literature review concerning research from the last two decades of the 20th century. The various studies contained within drew data from the Michigan Panel Study of Income Dynamics (PSID), the High School and Beyond Survey (HSB), the National Longitudinal Survey (NLS), and the National Longitudinal Survey of Youth (NLSY). They used OLS, logit, and probit models to analyze the intergenerational transmission of education. With simple regression analysis, these studies show that the correlation between parents' and their children's schooling is strong and robust when controlling for many variables, such as firstborn, socioeconomic status, resident region, and aspirations for their child, and using samples mostly from different countries in Europe and North America.

However, in more recent work, there has been some effort to distinguish causation from mere correlation in ability across generations (Carneiro, Meghir, & Parey, 2013; Holmlund et al., 2011). We first touch on three broad approaches that have been used to study the relationship of schooling between parents and their children: identical twins, adoptees, and instrumental variables. We then detail the limitations of these studies.

The different educational levels of identical twins, which were used to analyze the impact on their children, were the core of the identification strategy of the twins approach. Several studies have used twins data to discuss intergenerational effects of schooling. Behrman and Rosenzweig (2002) estimated the effect of parents' schooling on the schooling of their child using the data from the Minnesota Twin Registry (MTR), and concluded that the educational transmission effect of the father was about 0.35 and the mother's contributional counterpart was about −0.26. This means that each one-year increase in the father's total years of schooling contributed an additional 0.35 years of schooling to his child, while each additional year of the mother's total schooling decreased the child's schooling by 0.26 years. Antonovics and Goldberger (2005) used a subsample of the same MTR data and found a relatively high transmission effect of father-child, at about 0.48. However, some more recent studies pointed out that the transmission effect of parents on their child was much weaker. Bingley, Christensen, and Jensen (2009) estimated that the educational transmission effect of the father was only about 0.08 using a larger sample, but the mother-child transmission effect of education was not statistically significant. The transmission effect of education from the father and mother were about 0.16 and 0.10 respectively (Pronzato, 2012).

Using the adoption strategy for identification came from the idea that genetic transmission was absent between adopted children and their adoptive parents, so the results could capture all of the net educational transmission effect except for biological factors (Sacerdote, 2002). Plug (2004) estimated the effect of parents' schooling using the data from the Wisconsin Longitudinal Study 1992 (WLS), and pointed out that the transmission effect of the father was about 0.27 and the mother's contributional counterpart was about 0.28. Björklund, Lindahl, and Plug (2006) found much lower transmission effects, which showed that the father's effect was about 0.11 and the mother's effect was about 0.07.

Some studies estimated the transmission effect of education using educational reform as the instrumental variable (IV). The IV strategy relied on the idea that educational reform during parents' schooling had no direct influence on their children's schooling (Currie & Moretti, 2003). Chevalier (2004) estimated the transmission effect of education using the data drawn from the British Family Resources Survey 1994–2002 (BFRS), and found that the transmission effect of the father's was about 0.01 and the mother's was about 0.11. However, Black, Devereux, and Salvanes (2005) showed that the transmission effect of parents' schooling was not statistically significant.

The three types of studies above on the transmission of education—the twins approach, adoptee approach, and IV approach—-share gaps which we aim to fill. Principally, none concerned China, or even were restricted to rural areas. Most of them were studies from Europe and North America. Secondly, the use of cohorts was largely either absent or restricted to ten- to twenty-year periods. This makes the examination of time trends in education difficult. Thirdly, the sample sizes were quite small relative to ours, though a couple were similar in scope. Larger sample sizes are favorable for understanding the results' implications for the broader population. Finally, the representativeness of many of the studies, in terms of whether their samples reflect the population, is problematic for general applicability. In other words, the fact is that most people are neither twins nor adoptees, and further still, few belong to a group that witnesses significant educational reform during their schooling.

Though China is the most populous country in the world, it has attracted little attention in the literature. To our knowledge, there are only several studies which touch upon the intergenerational transmission of education in China. Sato and Li (2008) examined the intergenerational effects of family class origin on family members' education and found a drop caused by the class-based discrimination in the Maoist era and a rebound in the post-reform era. Yang and He (2015) used the Cultural Revolution as an instrument variable of the father's educational level and concluded that a 1-year increase in the father's schooling improved the probability of his child entering college by about 8 percentage points.

Both of these studies also have certain limitations upon which we would like to improve. They use data from the Chinese Household Income Project Series of 2002 (CHIPS 2002). As such, they share the limitation of using relatively older data compared to ours. Furthermore, just one of them focus on the rural population in China. Sato and Li (2008) use a cohort of those born in 1930–1982, while Yang and He (2015) use a cohort born in 1961–1983. Additionally, Sato and Li (2008) use a probit model to examine the binary outcome of the children achieving either more or < 10 years of education. This hampers the ability to understand the marginal contribution that each additional year of schooling of the parents has on their children. Yang and He (2015) have a similar limitation, as they only focus on whether the child entered college.

3. Data and identification strategy

3.1. Data

This study uses the dataset from the China Rural Development Survey (CRDS) collected by the Center for Chinese Agricultural Policy of the Chinese Academy of Sciences in April 2016. A multi-stage stratified cluster sampling procedure was used to select the sample. The sample provinces were randomly selected from each of China's major agro-ecological zones, not including Tibet, Hainan, and four province-level municipalities (Beijing, Tianjin, Shanghai, and Chongqing). Five sample counties were then selected from each province by a two-step procedure. First, the enumeration team listed all counties in each province in descending order of per capita gross value of industrial output (GVIO). GVIO was used based on the conclusions of Rozelle (1996) that GVIO is a good predictor of standard of living and development potential. He also concluded it is often more reliable than net rural per capita income statistics. Second, the five sample counties were selected randomly from each list. After the county selection was completed, the team then chose the sample townships and villages following the same procedure outlined above. Finally, the survey team combined village rosters and the survey team's own count of households, some of which were living in the village but not on the roster, to randomly choose 20 households in each village. Thus, a nearly nationally representative sample of 2026 households in 100 villages was selected.[1]

Our sampling strategy yields a sample that is not strictly nationally representative because it does not, for example, use census-based population counts in each survey year as sampling weights, but is nevertheless broadly reflective of China's rapidly changing rural population. Appendix Table 1 shows comparisons between our sample and data from China's National Bureau of Statistics (NBS), suggesting that they are reasonably similar.

To gauge how representative our sample is of the rest of rural China, we compare our CRDS findings to the data on rural households provided by the NBS. Their data for each statistic were gathered directly from the national census (NBS, 2016a, 2016b), except for average farmland, which was attained by dividing total farmland managed by rural households by the number of rural households. It is important to note the differences between our sample and the NBS, particularly pertaining to population size, sample size, and sampling method. For population, the census included sampling from every province, whereas ours excluded Tibet, Hainan, and four province-level municipalities (Beijing, Tianjin, Shanghai, and Chongqing). For sample size, the census in 2015 was not of the entire country, as in 2000 or 2010, but instead of around 1% of the total population, as in 2005. This is compared to our much smaller scale of only about 0.0013%. For sampling method, the census used a stratified, two-stage, probability proportion cluster sampling procedure, whereas ours was the aforementioned multistage stratified cluster sampling procedure.

As can be seen in the Appendix Table 1, our data for average age and years of schooling are quite similar to those of the NBS. The latter four statistical averages, however, vary more noticeably from the NBS data. Our sex ratio of males per hundred females is actually about 12.3% higher than the NBS, at 117.9 compared to 105. Household size, in terms of immediate family members per household, in our data is considerably higher, at 4.6 compared to 3.3. It is important to note that, while we have presented household size here using the same definition as the NBS to aid in clear comparison, we do not use this number for our empirical analyses later in the paper, in which we also include extended family members. Next, we feature nearly half the presence of minorities, at only 4.6% compared to 8.6%. We had anticipated this particularly large discrepancy in part due to our sample provinces (Jiangsu, Sichuan, Shaanxi, Jilin, and Hebei) not including the areas which feature large minority concentrations, such as the five autonomous regions in west and north China. Finally, our farmland, in terms of mu, where fifteen mu is equivalent to one hectare, is 7 mu compared to 5.7 mu.

In the rural households sampled for our study, we investigated all family members, including any children who have gone to school in other cities or provinces and those who have separated from the original household. In other words, we have investigated the original family and their extended family members. More importantly, this survey tracks at least three generations for each

[1] The data we used is from a follow-up investigation from 2005 to 2016, and the number of sample households was 2026 in the last round of the survey.

household: the parents, their children, and their grandchildren. As for the definition of children, it can be defined as follows: the direct biological off-spring of both parents who have finished their schooling and are > 16 years old at the time of sampling. It is possible that the current level of education is not the final level of education, and it should further be noted that the schooling of individuals in rural China rarely increases once the individual leaves school.

Those children, as defined above, may also have children of their own. Therefore, a family has two sets of distinct parent-child pairs: parent-child and child-grandchild. Instead of making a separate category for child-grandchild pairs, we simply relabel child-grandchild pairs as parent-child and pool them with the original parent-child pairs so as to conduct our analysis altogether. This allows us to take advantage of our survey and use FFE model. See Appendix fig. 3 for a common example. It briefly portrays intergenerational relationships within the family. In this example family, there are 5 pairs of parent-child relationships. They are: parent-child 1, parent-child 2, child1-grandchild 1, child1-grandchild 2, and child 2-grandchild 3. Variables such as parental schooling and the number of siblings may vary across individuals in different pairs of parent-child relationships; they would not be same for all the children of different generations in a family. For example, parent, child 1, and child 2 likely have different schooling years. Child 1, child 2, and grandchild 1–3 account for the entire sample of "children" in the study. Similarly, parent, child 1, and child 2 account for all sampled "parents" in the study. As there are several pairs of parents in a family, we can calculate our estimate in the case where there are parent-child relationships of more than two generations in a family. To our knowledge, there are no other studies that have collected similar detailed personal information about extended family members over time in rural China.

We collected the personal characteristics of all household members, such as gender, years of schooling, and birth year. In the survey, we coded individuals and relationships in the household in a way which enabled us to keep track of the different generations conveniently. Since we were only concerned with information for the schooling years of the immediate family, this means that information for the daughter-in-law, son-in-law, and so on, is absent. Finally, we obtain a sample which contains 6053 individuals and 2838 pairs of parents. Those individuals are 16 years old or above and have finished their schooling.

3.2. Identification strategy

In this paper, we adopt two methods of estimating the effect of parents' schooling on their children's schooling. These two methods are the OLS and the FFE models, which are discussed below in detail.

3.2.1. The OLS model

As the benchmark estimation, we use the OLS model firstly. For convenience, we define *Edu* as the total schooling years of an individual child. Edu_{dad}, Edu_{mom}, and Edu_{avg} represent the schooling years of the father, mother, and the average of the parents, respectively. Z is a vector of other factors which may influence the child's schooling years. These factors include: father's age when the child was born; mother's age when the child was born; number of the child's siblings who share the same mother and father; a set of dummy variable for male, non-Han, and firstborn, each with a value of 1 if the child has that characteristic and 0 otherwise; and a set of provincial dummies to capture regional factors that might affect the child's schooling years. We also add a set of birth cohort dummies to indicate into which birth cohort the child belongs. Specifically, we separate the sample into five cohorts indicating five separate decades: 1950–1959,[2] 1960–1969, 1970–1979, 1980–1989, and 1990–1999. This aims to help control the time trends in education, namely that schooling years generally increases for each successive cohort. Finally, we clustered the standard errors at the village level, allowing for correlation of households' influence on each other within the village.

Using Holmlund et al. (2011) and Pronzato (2012) for reference, we employ three empirical specifications to examine the intergenerational transmission of education from the father, mother, and parental average onto the child. For each child i in household h, we have:

$$Edu_{ih} = \alpha + \beta Edu_{dad_{ih}} + \rho Z_{ih} + \varepsilon_{ih} \tag{1}$$

$$Edu_{ih} = \alpha + \beta Edu_{mom_{ih}} + \rho Z_{ih} + \varepsilon_{ih} \tag{2}$$

$$Edu_{ih} = \alpha + \beta Edu_{avg_{ih}} + \rho Z_{ih} + \varepsilon_{ih} \tag{3}$$

where *Edu* is the schooling years of child. β is the coefficient of interest that captures the transmission of parents' schooling years to the child's. α is the constant term. ρ is the effect of other factors, including: father's age when the child was born, mother's age when the child was born, number of siblings, gender, non-Han, firstborn, provincial dummies, and birth cohort. ε is the error term. Compared with models (1) and (2), model (3) has the advantage that it controls for assortative mating, avoids multicollinearity, and produces more precisely estimated coefficients (Holmlund et al., 2011; Oreopoulos, Page, & Stevens, 2006).

3.2.2. The FFE model

A key issue facing the OLS model is potential endogeneity. In our case, this arises when parents and their children are linked by similar genetics and family culture. This connection between the generations may bias the results in a way which tends to overstate the effect of the parents' education on their children. In other words, should children follow a similar path to their parents which is

[2] In fact, we just have a few individuals who were born before 1950. Furthermore, after running some robustness checks, we found that data concerning the cohort of 1950–1959 and the years prior are quite similar. The empirical results show likewise there is no significant difference whether we include these individuals or not. Thus, we include all individuals born before the 1950s within that cohort itself.

statistically significant, the OLS model may describe this only as being a feature of intergenerational transmission, rather than as a combination alongside similar genetics and family culture.

In order to address the potential endogeneity as much as possible, we use the FFE model to estimate the transmission effect of parents' schooling years on their children's schooling years while eliminating these other contributing factors which have effect on both parents and their children. We did so by taking advantage of our survey in which we gathered data on three generations within the same family. This allows us to make different pairings of each parent with each child within the same family. The educational attainment of those parents in the same family usually varies from each other. In the interest of having a further contribution to the literature, we have not seen any other studies on China specifically which employ the FFE to address the endogeneity problem. Despite these advantages, there still remain some disadvantages to using the FFE. Particularly, the model assumes that children within the same family have identical family culture and genetics. Clearly, this is not strictly the case.

In the FFE model, for each child i in household h, we have:

$$Edu_{ih} = \alpha + \beta Edu_{dad_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{4}$$

$$Edu_{ih} = \alpha + \beta Edu_{mom_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{5}$$

$$Edu_{ih} = \alpha + \beta Edu_{avg_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{6}$$

The definitions of α and β are the same as above. ε is the error term. As mentioned, a pooled regression such as OLS ignores the unobservable characteristics ν_h shared in each household, like genetics, family culture, and so on, which influence the education both of parents and of children. By using FFE, we can eliminate ν_h from the equation by differencing the equation above in the following way:

$$Edu_{ih} - \overline{Edu_h} = \beta(Edu_{dad_{ih}} - \overline{Edu_{dad_h}}) + \rho(Z_{ih} - \overline{Z_h}) + (\nu_h - \nu_h) + (\varepsilon_{ih} - \overline{\varepsilon_h}) \tag{7}$$

$$Edu_{ih} - \overline{Edu_h} = \beta(Edu_{mom_{ih}} - \overline{Edu_{mom_h}}) + \rho(Z_{ih} - \overline{Z_h}) + (\nu_h - \nu_h) + (\varepsilon_{ih} - \overline{\varepsilon_h}) \tag{8}$$

$$Edu_{ih} - \overline{Edu_h} = \beta(Edu_{avg_{ih}} - \overline{Edu_{avg_h}}) + \rho(Z_{ih} - \overline{Z_h}) + (\nu_h - \nu_h) + (\varepsilon_{ih} - \overline{\varepsilon_h}) \tag{9}$$

where "-" indicate the average of each variable in each family. In this way, we can eliminate the unobservable characteristics as much as possible and partially address the endogenous problem.

4. Results

4.1. Descriptive statistics of children and their parent's schooling

Table 1 shows the schooling years of parents and their children. According to our data, the average schooling of the children is 9.07 years (row 6, column 2). The father's average schooling is 6.10 years (row 6, column 5). The average schooling of the mother is 3.76 years (row 6, column 8). Thus, the average schooling years of father is about 1.6 times the schooling years of the mother in rural families.

Considering the heterogeneity of different birth years, we group the data by children birth cohort every 10 years. Trends among cohorts are the basis for our observations of educational change over time. The first cohort is the sample born before the year of 1960, which contains 201 people. The average schooling of these members is 6.47 years, with their father's and mother's average schooling being 2.04 years and 0.67 year respectively (row 1, columns 2, 5, and 8). Individuals born between 1960 and 1969 have 7.51 schooling years on average, which is equivalent to the first year of middle school. However, the average schoolings of the father and mother for this cohort are 4.24 years and 1.98 years respectively, which means they did not complete primary school (row 2, columns 2, 5, and 8).

Individuals in the third cohort have typically attended second grade in middle school, as the average schooling is 8.22 years. Their parents still did not complete primary school; the average schoolings of father and mother are 5.50 years and 2.73 years respectively

Table 1
Description of child's schooling by birth cohorts in rural China.

	Child's birth cohort	Child			Father			Mother			Parents' average		
		Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
(1)	1950s (1950–1959)	201	6.47	3.92	119	2.04	3.02	119	0.67	2.37	119	1.36	2.11
(2)	1960s (1960–1969)	1081	7.51	3.39	509	4.24	3.95	509	1.98	3.00	509	3.11	2.92
(3)	1970s (1970–1979)	1739	8.22	3.28	992	5.50	3.51	992	2.73	3.19	992	4.12	2.73
(4)	1980s (1980–1989)	1831	10.14	3.70	1229	7.37	3.49	1229	4.80	4.09	1229	6.09	3.18
(5)	1990s (1990–1999)	1201	10.49	3.24	1031	7.48	3.02	1031	5.82	3.47	1031	6.65	2.67
(6)	Total	6053	9.07	3.67	2838	6.10	3.78	2838	3.76	3.81	2838	4.93	3.24

Data source: Authors' survey.

Table 2
Relationship between parent' and child's schooling: estimations of the OLS model.

	Explanatory variables	Schooling of child (years)								
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
(1)	Father's schooling	0.374***	0.270***	0.262***						
		(0.016)	(0.018)	(0.019)						
(2)	Mother's schooling				0.367***	0.267***	0.255***			
					(0.019)	(0.020)	(0.019)			
(3)	Average schooling of both parents							0.510***	0.401***	0.396***
								(0.018)	(0.021)	(0.021)
(4)	Father's age when the child was born			−0.001			−0.036*			−0.001
				(0.018)			(0.021)			(0.019)
(5)	Mother's age when the child was born			0.077***			0.106***			0.083***
				(0.019)			(0.020)			(0.019)
(6)	Firstborn (1 = yes)			0.526***			0.436***			0.462***
				(0.118)			(0.113)			(0.112)
(7)	Number of siblings			−0.279***			−0.246***			−0.209***
				(0.069)			(0.065)			(0.064)
(8)	Male (1 = yes)			0.475***			0.491***			0.515***
				(0.104)			(0.108)			(0.105)
(9)	Non-Han (1 = yes)			−0.300			−0.255			−0.135
				(0.291)			(0.304)			(0.282)
(10)	1960s		0.398	0.074		0.551*	0.271		0.251	−0.000
			(0.310)	(0.318)		(0.321)	(0.321)		(0.312)	(0.317)
(11)	1970s		0.737**	0.119		1.017***	0.462		0.525	0.039
			(0.333)	(0.364)		(0.352)	(0.369)		(0.339)	(0.365)
(12)	1980s		2.198***	1.308***		2.412***	1.621***		1.679***	1.000**
			(0.340)	(0.401)		(0.363)	(0.404)		(0.351)	(0.400)
(13)	1990s		2.527***	1.512***		2.512***	1.611***		1.813***	1.047***
			(0.335)	(0.383)		(0.356)	(0.394)		(0.346)	(0.389)
(14)	Province dummies		Included	Included		Included	Included		Included	Included
(15)	Constant	6.773***	6.562***	5.701***	7.687***	7.307***	6.503***	6.545***	6.650***	5.165***
		(0.141)	(0.484)	(0.685)	(0.131)	(0.416)	(0.624)	(0.144)	(0.434)	(0.635)
(16)	Observations	6053	6053	6053	6053	6053	6053	6053	6053	6053
(17)	R-squared	0.148	0.203	0.223	0.145	0.202	0.220	0.201	0.233	0.250
(18)	Number of households	1915	1915	1915	1915	1915	1915	1915	1915	1915

Data source: Authors' survey. Notes: Robust standard errors in parentheses.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

(row 3, columns 2, 5 and 8). The last two cohorts contain the sample of individuals who were born between 1980 and 1999, marking the era of reform and opening-up. The average schooling of these two cohorts is above 10 years, which means they attended senior high school. On average, their father had completed primary school, at about 7.4 years, and their mother neared graduation from primary school, at about 5.3 years (rows 4 and 5, columns 2, 5, and 8).

Table 1 also shows that average schooling attainment grew from 6.47 years for the oldest cohort to 10.49 years for the youngest cohort (rows 1 and 5, column 2). Each group's schooling is much higher than their parents, regardless of birth cohort.

4.2. Results of the OLS model

According to the OLS estimation, there is a positive relationship between parents' schooling and their children's schooling (Table 2). The estimated effects of the father's schooling show—consistently with previous studies—that more educated fathers increase their children's schooling. We find that the impact of the father-child transmission effect is 0.374 ($P < 0.01$), which means that each 1-year increase in the father's schooling will on average lead to about a 0.374-year increase in their children's schooling (row 1, column 1). When we control for time trends and regional factors, this effect reduces to 0.270 ($P < 0.01$) (row 1, column 2). When we further control for personal and parental characteristics, we find the magnitude of the transmission effect is almost unchanged at 0.262 ($P < 0.01$) (row 1, column 3). The result is not obviously higher compared to other studies (Antonovics & Goldberger, 2005; Behrman & Rosenzweig, 2002). The mother's schooling years almost has the same effect on their children's schooling years as the father's.

However, the impact of parents' average schooling on children's schooling is positive and much stronger than the father's or mother's alone. The result shows that each 1-year increase of parents' education will on average add about 0.510 ($P < .01$) years to their children's education (row 3, column 7). When we further control for other covariates, the transmission effect is reduced to 0.396 ($P < 0.01$) (row 3, column 9).

Though the results of our study show that average schooling of both parents plays a stronger role than the father's schooling or the mother's schooling alone, the coefficient of the average schooling of both parents' is not comparable to the coefficients of the father's schooling or the mother's schooling individually. The results need to be interpreted cautiously. The average schooling of both parents increasing by one year is equivalent to both the father's schooling and the mother's schooling increasing by one year. Thus, the average schooling of both parents plays a stronger role than either the father's schooling or the mother's schooling, but it is still at most similar to the summed effects of the father's and the mother's. This is consistent with the previous studies (Björklund et al., 2006; Holmlund et al., 2011; Oreopoulos et al., 2006), which all show that the effect of both parents' schooling is less than or similar to the summed effects of father's schooling and the mother's schooling. For example, Holmlund et al. (2011) shows that the intergenerational transmission effect of father-to-children is 0.15, the mother's is 0.20, and the parents' combined is 0.34.

Table 2 also displays that children's schooling is correlated with other characteristics. The father' age when a child is born has almost no effect on his children's schooling (row 4, columns 3, 6, and 9). However, the mother' age when a child is born has a positive effect on her children's schooling. On average, if a mother delays birth for one additional year, her children's schooling years increase by about 0.077–0.106 years ($P < 0.01$) (row 5, columns 3, 6 and 9). The first child in the household has an advantage in achieving education of about 0.5 years ($P < 0.01$) (row 6, columns 3, 6, and 9). However, the number of siblings has a negative effect on the children's educational attainment by about 0.209–0.279 years ($P < 0.01$) (row 7, columns 3, 6, and 9). When other conditions do not change, male children's schooling on average is about 0.5 years more than females ($P < 0.01$) (row 8, columns 3, 6, and 9). Meanwhile, the results show that there is no significant difference between Han and minority nationalities in achieving education (row 9, columns 3, 6, and 9). As the decades progress, almost every successive cohort achieves more education than the preceding cohort (rows 10–13, columns 3, 6, and 9).

4.3. Results of the FFE model

By eliminating the influence of genetics and family culture, we can get a more precise estimation of the transmission effect of education across two generations (Table 3). These results are consistent with those of the OLS model, a positive relationship between parents' schooling years and their children's schooling years. When just using a simple univariate specification that focuses on the intergenerational transmission of education, the estimation results show that a 1-year increase of the father's schooling will, on average, result in a 0.389-year increase of their children's schooling ($P < 0.01$) (row 1, column 1). When we further control for the impact of time trends, regional factors, and personal and parental characteristics, the transmission effect reduces to about 0.081 ($P < 0.01$) (row 1, column 3). The results are relatively low compared to those of previous studies (Behrman & Rosenzweig, 2002; Björklund et al., 2006; Plug, 2004; Pronzato, 2012). The mother's schooling has almost the same effect on their children's schooling as the father's.

Similar to the results of OLS model, the effect of parents' average schooling years on their children's schooling years is also larger than that of the father's or mother's alone, from about 0.08 to 0.13 ($P < 0.01$) (row 1, column 3; row 2, column 6; row 3, column 9). Due to the partial alleviation of the endogenous problem, the transmission effect undergoes a large reduction. The father-child and mother-child transmission effects reduce from about 0.26 in the OLS model to 0.08 in the FFE model ($P < 0.01$) (row 1, column 3; row 2, column 6).

Table 3 also reports the effects of other characteristics on the children's schooling. For example, the father' age and mother's age at which a child is born has no effect on their children's schooling (rows 4 and 5, columns 3, 6, and 9), which is different from the result of the OLS model. The effect of whether the child is the first child of their parents estimated by the FFE model is similar to the effect in the results of the OLS model. The effect of the number of siblings estimated by the FFE model is also similar to the results of the OLS model, about 0.4 years on average (row 6, column 3, 6, and 9). Male children's schooling is about 0.73 years more than females, which is much bigger than that of the OLS model ($P < 0.01$) (row 8, columns 3, 6, and 9). They both imply that preference for a son exists in the educational choices within the average household.

4.4. Further tests and robustness checks

4.4.1. Lowess curve

Most of the existing literature finds that the intergenerational transmission of education has a linear relationship (Holmlund et al., 2011). Therefore, OLS is employed where the schooling of the child is regressed on the schoolings of the parents (Hertz et al., 2007; Holmlund et al., 2011; Pronzato, 2012). Furthermore, when we run a locally weighted scatterplot smoothing (Lowess) with our data, the results also show an almost linear relationship between the schoolings of parents and that of their child (Appendix Figs. 1 and 2).

4.4.2. Nonlinear robustness check

As a robustness check, we also include a quadratic term for the father's, mother's, and parents' average schooling years. This simply means we square the value for education in each of our previous equations. In other words, our robustness check equations take the form:

$$Edu_{ih} = \alpha + \beta Edu_{dad_{ih}} + \gamma Edu^{2}{}_{dad_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{10}$$

$$Edu_{ih} = \alpha + \beta Edu_{mom_{ih}} + \gamma Edu^{2}{}_{mom_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{11}$$

$$Edu_{ih} = \alpha + \beta Edu_{avg_{ih}} + \gamma Edu^{2}{}_{avg_{ih}} + \rho Z_{ih} + \nu_h + \varepsilon_{ih} \tag{12}$$

Table 3
Relationship of parent' and child's schooling: estimations of the FFE model.

	Explanatory variables	Schooling of child (years)								
		(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
(1)	Father's schooling	0.389***,**,*	0.112***	0.081***						
		(0.032)	(0.023)	(0.023)						
(2)	Mother's schooling				0.446***	0.120***	0.076***			
					(0.024)	(0.026)	(0.025)			
(3)	Average schooling of both parents							0.553***	0.175***	0.125***
								(0.030)	(0.032)	(0.031)
(4)	Father's age when the child was born			0.024			0.021			0.028
				(0.029)			(0.030)			(0.029)
(5)	Mother's age when the child was born			0.010			0.014			0.010
				(0.031)			(0.031)			(0.031)
(6)	Firstborn (1 = yes)			0.422***			0.412***			0.412***
				(0.122)			(0.123)			(0.122)
(7)	Number of siblings			−0.410***			−0.409***			−0.391***
				(0.092)			(0.094)			(0.095)
(8)	Male (1 = yes)			0.729***			0.728***			0.731***
				(0.109)			(0.109)			(0.109)
(9)	Non-Han (1 = yes)			−0.082			−0.097			−0.083
				(0.723)			(0.730)			(0.721)
(10)	1960s		1.263***	1.022***		1.284***	1.036***		1.226***	0.975***
			(0.252)	(0.240)		(0.252)	(0.239)		(0.252)	(0.240)
(11)	1970s		1.929***	1.497***		1.931***	1.503***		1.836***	1.408***
			(0.261)	(0.283)		(0.264)	(0.280)		(0.262)	(0.282)
(12)	1980s		4.124***	3.195***		4.011***	3.145***		3.868***	3.023***
			(0.294)	(0.307)		(0.308)	(0.305)		(0.304)	(0.306)
(13)	1990s		4.556***	3.408***		4.392***	3.333***		4.225***	3.195***
			(0.323)	(0.318)		(0.343)	(0.316)		(0.345)	(0.323)
(14)	Constant	6.681***	5.450***	5.670***	7.390***	5.748***	5.896***	6.330***	5.449***	5.533***
		(0.199)	(0.278)	(0.500)	(0.090)	(0.252)	(0.489)	(0.146)	(0.273)	(0.490)
(15)	Observations	6053	6053	6053	6053	6053	6053	6053	6053	6053
(16)	R-squared	0.069	0.196	0.223	0.109	0.196	0.222	0.117	0.198	0.224
(17)	Number of households	1915	1915	1915	1915	1915	1915	1915	1915	1915

Data source: Authors' survey.
Notes: Robust standard errors in parentheses.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Appendix Table 2 presents the outcome of the OLS regression in addition to our robustness checks. For the sake of brevity, we only present the key variables from before (i.e., education of parents) and the new variables described above. The results show that all the coefficients of variables in the robustness check models are consistent with the benchmark model. Moreover, when we put the quadratic term for father's, mother's, and parents' schooling years into the equation, we find that the transmission effect decreases only slightly. However, most of the newly added quadratic term coefficients are not significant.

We also test two dummy variables indicating whether the father, mother, or both parents attended senior high school besides just looking at their total schooling years. A value of one for the father's new dummy variable indicates that he didn't attend senior high school, and zero otherwise. This is the same for the mother's. A value of one for both indicates that the pair didn't attended high school, and zero for either indicates that the pair did. The addition of these dummy variables therefore gives us three more equations for our robustness checks: one each for the father's, mother's, and combined parents' status of having attended or not attended senior high school.

When we add only the dummy variable for whether the father attended senior high school, it is significant at the 10% level (Appendix Table 2, row 4, column 3). We find that the father's transmission of education for schooling years remains similar to the results of the benchmark OLS model (row 1, columns 1 and 3). The mother's dummy variable, however, is not significant at the 10% level (row 5, column 6), and the transmission of education for schooling years is similar to the benchmark OLS model (row 2, columns 4 and 6). The parents' combined transmission of education on the child is much larger than the father's (row 3, columns 7–9). All three transmission of education coefficients—the father's, mother's, and parents' combined—are significant at the 1% level (row 1, columns 1–3; row 2, columns 4–6; row 3, columns 7–9).

Appendix Table 3 presents the outcome of the FFE model as presented in the paper as well as the robustness check above. Similar to Appendix Table 2, we only present the key variables and the newly added variables. The results from the robustness checks show that all the coefficients of the key variables that we are interested in are consistent with the benchmark model. As before, most coefficients of the new added variables are not significant. Appendix Tables 2 and 3 show that the empirical results in Section 4.2 and 4.3 are reliable, and furthermore, that our original results are robust.

4.4.3. Robustness check for village income level of each cohort

Some literature shows that an individual's educational level is correlated with their socioeconomic status, part of which may be represented by household assets and income (Gong et al., 2012; Yang & Qiu, 2016). To the best of our knowledge, there is little to no data for rural China which tracks household assets or income from thirty years ago, and certainly not for five or six decades ago. Due to the data limitation, we cannot control the effect of household assets and household income on the schooling of sampled individuals. Certainly, this is one main limitation of this study. We employ a regression to control for the influence of income as much as possible. Furthermore, we account for the economic status of different villages at different times for each birth cohort. As there are 100 villages and 5 different birth cohort dummies, this results in 400 new dummy variables (100 × 4). As with our original model, a value of 0 for all birth cohort dummies indicates membership to the fifth cohort. We simply include these dummies in our Z variable and run the model again. The outcomes are presented in the Appendix Tables 4 and 5.

Appendix Table 4 shows that the transmission effect almost has no change compared to the OLS model used before, though the transmission effect has decreased for all three of father-child, mother-child, and parent-child (row 1, columns 1 and 2; row 2, columns 3 and 4; row 3, columns 5 and 6). Consistent with the FFE model used before, Appendix Table 5 shows that the transmission effect exists (row 1, columns 1 and 2; row 2, columns 3 and 4; row 3, columns 5 and 6). Both tables show that the results still are robust when the income level has been considered.

4.4.4. Robustness check for parental migration to gain employment

We were only able to collect information about whether individuals migrated for employment starting from the year 1998 onward. This subsample, for which we may include data concerning employment gained by migration, therefore contains 2078 individuals. Employment gained by migration of the father is expressed as a dummy variable, equal to 1 if the father has at least one year of employment by migration during the child's schooling years, and 0 if not. A second dummy variable is created for the mother using the same method. Adding these two variables to our equations from Section 3.2, we perform the same OLS estimation on this subsample to compare their results with our benchmark OLS estimation. Shown in Appendix Table 6, the results of this robustness check suggest that the findings of the OLS model used earlier in the paper and those of this new OLS model are similar. We cannot take advantage of the FFE due to the lack of parent-child pairs.

Appendix Table 6 shows the results. It shows that the transmission effect is similar compared to the OLS model used before (row 1, columns 1 and 2; row 2, columns 3 and 4; row 3, columns 5 and 6). The transmission effect of the father's, mother's, and parents' average schooling are 0.294, 0.249 and 0.413 respectively (row 1, column 2; row 2, column 4; row 3, column 6). These are near to the results of the benchmark OLS model, which are 0.262, 0.255 and 0.396. The migration experience of the father and mother are not significant even at the10% level.

4.4.5. Test the influence of parents' schooling gap

Appendix table 7 presents the general picture of the extent to which the schooling of a couple differs. Similar to the above tests and robust checks, we only present the key variables and the newly added variables. As shown in this table, according to our data, 32.23% of the sampled' fathers and mothers have the same schooling years (row 1, column 1). Those couples with a schooling gap that is between 1 year and 3 years comprise 28.21% (rows 2–4, column 1). Meanwhile, 84% of couples have a schooling gap of < 6 years (rows 1–7, column 1).

Appendix table 8 presents the results of a subgroup analysis by using the OLS and FFE methods. Subgroups are separated based on the absolute value of schooling gap within a couple. Though the results of the OLS model show some minor differences among different subgroups, the results of FFE model show that the roles of the average schooling of both parents are different between that of the couples with similar years of schooling and that of those with significantly different schooling. The results show that the intergenerational transmission effect of parents to their children functions well when both parents have the similar schooling years.

4.4.6. Test the regional difference

We have also tested if there is significant difference across regions for the role of parents' education, besides just by including provincial dummies. The results are presented in Appendix table 9. As shown in the table, there are two provinces which seem to have a lower intergenerational transmission effect than the other provinces (rows 6 and 7, column 4; rows 10 and 11, column 5; rows 14 and 15, column 6). These two provinces are Jilin and Hebei. However, the interpretation of how this result occurred requires further research and we have not found related discussions in the previous studies.

4.5. Comparison between before and after the OCP and reforms

To characterize the evolution of the transmission effect of education between two generations before and after China's major reforms, we make a subgroup analysis. The original sample is divided into two subgroups which are separated by the birth year of 1980. The first subgroup includes all those born through the end of 1980, and the second subgroup includes all those born in 1981 onward. In September of 1980, China's One-Child Policy (OCP) was formally implemented, decreeing that all families may have no more than one child (Yang, 2013). It was also around this time that China began its milestone era of reform and the opening-up policies which mostly refer to the program of economic reforms termed "Socialism with Chinese characteristics". We aim to capture the effects of these two dramatic changes to China's social fabric on our model. We run the OLS and FFE models again, taking out the previous birth cohorts and using these two subgroups instead. Using the same logic as before, we focus primarily on the results of the FFE model because we believe them to be more precise in not overstating the effect of the intergenerational transmission of education.

Table 4
Relationship of father's and child's schooling: before and after the OCP and reforms.

Explanatory variables		Before the end of 1980		After the start of 1981	
		OLS (1)	FFE (2)	OLS (3)	FFE (4)
(1)	Father's schooling	0.235***	−0.041	0.307***	0.268***
		(0.025)	(0.188)	(0.027)	(0.051)
(2)	Father's age when the child was born	0.015	0.469*	−0.057**	−0.063
		(0.024)	(0.241)	(0.028)	(0.077)
(3)	Mother's age when the child was born	0.070***	−0.364	0.110***	0.151**
		(0.025)	(0.241)	(0.029)	(0.060)
(4)	Firstborn (1 = yes)	0.367**	0.410**	0.399**	0.331
		(0.145)	(0.157)	(0.177)	(0.237)
(5)	Number of siblings	−0.261***	−1.205***	−0.564***	−0.512***
		(0.064)	(0.451)	(0.123)	(0.163)
(6)	Male (1 = yes)	1.040***	1.374***	−0.229	0.058
		(0.154)	(0.160)	(0.139)	(0.161)
(7)	Non-Han (1 = yes)	0.641	1.144*	−0.222	−0.460
		(0.458)	(0.644)	(0.502)	(0.658)
(8)	Province dummies	Included	Included	Included	Included
(9)	Constant	5.140***	6.665***	7.816***	6.697***
		(0.572)	(1.725)	(0.927)	(1.511)
(10)	Observations	3161	3161	2892	2892
(11)	R-squared	0.152	0.091	0.140	0.039
(12)	Number of households	1155	1155	1537	1537

Data source: Authors' survey.
Note: Robust standard errors in parentheses.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

The FFE results indicate that the impact of the father's schooling years on his children's is not statistically significant even at a 10% significance level for the subgroup born before the OCP and reforms era (Table 4, row 1, column 2). As for the subgroup born from 1981 onward, the transmission effect of education is statistically significant, and the coefficient is 0.268 (row 1, column 4). For individuals born before the OCP and reforms era, male children's schooling is about 1.374 years more than female children's (row 6, column 2). But for the individuals who were born in the year of 1980 onward, the education gap between male children and female children is not statistically significant (row 6, column 4). Counterintuitively, being a male only has a significant positive effect on the years of schooling before the OCP and opening-up, but not after. One might expect the opposite, as the country's sudden and strong preference for boys during this time period is well-known. This is likely due to the compulsory education reforms in 1986, which ensured that girls had the same educational opportunities as boys. The effect of whether the child is the first child of their parents and the effect of the sibling number decreased compared to the subgroup born before the era of the OCP and reforms (row 4, columns 2 and 4; row 5, columns 2 and 4). The mother-child estimation results of the transmission effect are similar to those of the father-child, but the extent of the effect is smaller (Table 5, row 1, column 4).

The parent-child estimation results of the transmission effect are presented in Table 6. The OLS estimation shows that a one-year increase in the average parental schooling will lead to about a 0.343-year increase in the children's schooling if they were born before 1980 (row 1, column 1). The FFE estimation shows that the impact of parents' schooling on their children's is not statistically significant even at a 10% significance level (row 1, column 2). However, the OLS and FFE estimations both show a statistically significant transmission effect in the subgroup born after the 1980s. The OLS estimation indicates that a one-year increase of parents' schooling will lead to about 0.441-year increase in the children's schooling years (row 1, column 3). The FFE estimation likewise shows a statistically significant transmission effect between parents and children (row 1, column 4).

The education achievement gap between the subgroups before and after the 1980s can be seen clearly. Some possible explanations could be described as below. First, in the era of 1950s and before, the educational level of rural Chinese residents was universally low, and the proportion of illiteracy was very high. Second, the implementation of the compulsory education law in 1986, which popularized primary and secondary education, has led to a significant increase in the level of schooling of the rural population, which has narrowed the differences in schooling among the population born in those days.

Following the Cultural Revolution in 1966, the vast majority of intellectuals became the object of persecution, and the education system was, to a large extent, destroyed. As a result, the public grew to have bad impressions of those who attained a high level of education. Moreover, following 1966, large numbers of schools and colleges could no longer enroll students. This also likely hindered the intergenerational transmission of education. In general, a series of educational policies and political movements in the Mao era probably weakened the relationship between parents' education and their children's education.

Meanwhile, that the schooling gaps between men and women have narrowed can be inferred from these results. The reasons may be as follows. First, the implementation of the nine-year compulsory education policy, which is not based on gender, raised the

Table 5
Relationship of mother's and child's schooling: before and after the OCP and reforms.

		Before the end of 1980		After the start of 1981	
Explanatory variables		OLS (1)	FFE (2)	OLS (3)	FFE (4)
(1)	Mother's schooling	0.203***	0.120	0.285***	0.166***
		(0.026)	(0.195)	(0.022)	(0.050)
(2)	Father's age when the child was born	−0.026	0.464*	−0.074***	−0.036
		(0.026)	(0.248)	(0.027)	(0.085)
(3)	Mother's age when the child was born	0.101***	−0.359	0.138***	0.121*
		(0.027)	(0.248)	(0.027)	(0.066)
(4)	Firstborn (1 = yes)	0.319**	0.407**	0.349**	0.322
		(0.148)	(0.157)	(0.173)	(0.236)
(5)	Number of siblings	−0.302***	−1.081**	−0.353***	−0.548***
		(0.065)	(0.532)	(0.114)	(0.155)
(6)	Male (1 = yes)	1.034***	1.374***	−0.173	0.028
		(0.156)	(0.159)	(0.142)	(0.169)
(7)	Non-Han (1 = yes)	0.488	1.144*	0.031	−0.561
		(0.449)	(0.644)	(0.570)	(0.691)
(8)	Province dummies	Included	Included	Included	Included
(9)	Constant	6.590***	5.860***	8.413***	7.892***
		(0.512)	(1.911)	(0.843)	(1.384)
(10)	Observations	3161	3161	2892	2892
(11)	R-squared	0.126	0.091	0.149	0.024
(12)	Number of household	1155	1155	1537	1537

Data source: Authors' survey.
Note: Robust standard errors in parentheses.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

Table 6
Relationship of parents' and child's schooling: before and after the OCP and reforms.

		Before the end of 1980		After the start of 1981	
Explanatory variables		OLS (1)	FFE (2)	OLS (3)	FFE (4)
(1)	Average schooling of both parents	0.343***	0.009	0.441***	0.360***
		(0.027)	(0.215)	(0.029)	(0.073)
(2)	Father's age when the child was born	0.012	0.469*	−0.051*	−0.050
		(0.024)	(0.244)	(0.026)	(0.080)
(3)	Mother's age when the child was born	0.078***	−0.364	0.113***	0.143**
		(0.025)	(0.245)	(0.026)	(0.063)
(4)	Firstborn (1 = yes)	0.383***	0.409**	0.325*	0.331
		(0.144)	(0.157)	(0.171)	(0.234)
(5)	Number of siblings	−0.225***	−1.181**	−0.338***	−0.410**
		(0.064)	(0.476)	(0.111)	(0.169)
(6)	Male (1 = yes)	1.070***	1.374***	−0.151	0.054
		(0.152)	(0.159)	(0.142)	(0.162)
(7)	Non-Han (1 = yes)	0.406	1.143*	−0.199	−0.492
		(0.385)	(0.644)	(0.584)	(0.683)
(8)	Province dummies	Included	Included	Included	Included
(9)	Constant	4.891***	6.373***	6.887***	6.107***
		(0.540)	(1.823)	(0.860)	(1.552)
(10)	Observations	3161	3161	2892	2892
(11)	R-squared	0.162	0.091	0.184	0.040
(12)	Number of households	1155	1155	1537	1537

Data source: Authors' survey.
Note: Robust standard errors in parentheses.
*** $p < 0.01$.
** $p < 0.05$.
* $p < 0.1$.

educational level of women. Second, women in rural areas have had only a restricted number of employment channels. Since there existed, and still exits, gender discrimination in the labor market in these areas, it also prompted them to improve their level of education and therefore enhance their competitiveness in the labor market. Doing so might also give them the option of migrating to areas with less gender discrimination.

Furthermore, we also get an obvious contrast of the transmission effect when we compare the results of the subgroup estimation. The intergenerational transmission effect of education is not statistically significant before China's OCP and reforms, yet becomes significant afterward. This may imply that parents became better able to pass down their own education to their children following the OCP and reforms. However, this could also mean that parents simply came to recognize the importance of education for their children within the rapidly changing society, and thus push these new values rather than become better at passing down their own education. Nevertheless, the results suggest that actions taken by the government to promote schooling for individuals, especially those who were low-educated, were well-considered, and achieved their desired effect on the population.

5. Conclusion and discussion

In this paper, we have estimated the intergenerational transmission effect of education in rural China by using samples covering > 2000 households in 100 villages. By using the ordinary least squares and family fixed effect models, we find evidence of close correlation between the father's, mother's, and combined parents' schooling years and their children's schooling years. Our results show that the OLS model overestimates the transmission effect of parents' education on children's education compared to the FFE model, which matches our predictions. By using the latter model, we get a relatively smaller but statistically significant transmission effect of parents' schooling on their children's schooling. We believe the results of this model to be a more accurate estimation of the transmission effect because they work toward eliminating unobservable family characteristics which may also affect children's schooling years. Notably, the results stand up to several different robustness checks which each aim to account for a variety of different factors.

There are some limitations to our study. Though we attempted to control for them using the FFE model, there are unobservable family characteristics which may still push toward an overstatement of the intergenerational transmission effect. We also attempted to deal with other factors, such as village income level of each cohort, which may contribute to the children's education continuing down the same path trodden by their parents. We have not run checks for things such as social attitudes toward a given family, which may pressure members of that family toward certain levels of education. These types of factors are problematic to deal with from a data perspective to begin with, and it is difficult to estimate the extent to which they affect the education of children.

Nonetheless, the results from our data suggest that the policy implications for the intergenerational transmission of education, specifically in rural areas of China, are undeniable. Education has positive effects on individuals' access to off-farm employment opportunities, and improves their health status, productivity, and wealth (De Brauw, Huang, Rozelle, Zhang, & Zhang, 2002; Hertz et al., 2007). In the short-term, it is evident from our study that raising the bar for education of one generation has historically led to statistically significant educational gains for the next. As such, policies aimed at improving the educational attainment of young people in rural areas should be implemented by government. Among these actions may be measures which expand opportunities for access to senior high school and college. The government should also consider programs which help to loosen the financial fetters felt by households in the short-term when they invest in their family's education. As our study suggests, these policies may have significant effects on the improvement of young people's educational attainment.

Moreover, the empirical results of this study have broader policy implications down the road. It is implied that, if the results of our paper are to hold for the future, increasing the amount of schooling for all rural children today would lead to an increase in the schooling of their next generation. Thus, the policies stated above have long-term effects for the accumulation of educational attainment in each household. By taking advantage of the natural effect of the intergenerational transmission of education from parents to their children on a national scale, the government may ensure long-term fortitude among its rural people. In this way, China can promote increasing gains for its acquisition of human capital, and tap into this foundation for sustainable growth and development.

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Appendix A. Supplementary data

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