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Compensating for unequal parental investments in schooling

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Abstract This paper investigates how rural families in China use marital and post-marital transfers to compensate their sons for unequal schooling expenditures. Using a common behavioral framework, we derive two methods for estimating the relationship between parental transfers and schooling investments: the log-linear and multiplicative household fixed-effects regression models. Using data from a unique household-level survey, we strongly reject the log-linear specification. Results from the multiplicative model suggest that when a son receives 1 yuan less in schooling investment than his brother, he obtains 0.47 yuan more in transfers as partial compensation. Since our measure of transfers represents a substantial fraction of total parental transfers, sons with more schooling likely enjoy higher lifetime consumption. Redistribution within the household may be limited by either the parents' desire for consumption equality or bargaining constraints imposed by their children. Controlling for unobserved household heterogeneity and a fuller accounting of lifetime transfers are quantitatively important.

Keywords Household model · Parental investment · Marriage market · Transfers

JEL Classification D13 · J12 · J13

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1 Introduction

Depending on their children's abilities and labor market opportunities, parents may choose different schooling investments for their children. After schooling investments are made, parents can use inter vivos transfers to mitigate consumption inequality among siblings. Becker's (1991) unitary model of the family predicts that altruistic parents use monetary transfers to fully compensate children for resulting consumption inequalities. Bargaining and exchange models of intra-household allocations, however, predict that parents may not fully compensate children for the resulting inequalities. Thus, how much parents compensate children for unequal parental investments in schooling is an empirical question.

The empirical literature has examined three important dimensions of the issue. First, it shows that parents invest more in the schooling of those children for whom returns are higher due to either higher abilities or better labor market opportunities relative to their siblings.[1] Second, several studies use panel data to study the extent of intra-family insurance against exogenous high frequency (annual) income or wage shocks of family members.[2] They conclude that family insurance is quantitatively small relative to the predictions of the unitary model. Furthermore, many parents do not give annual monetary transfers to their children. And third, several studies show that parents occasionally provide large monetary gifts to their children such as wedding presents and help with down payments for housing purchases.[3] Both the incidence and size of these gifts are inversely related to the permanent incomes of their children. Parents also give bequests.[4] Unlike family insurance against high frequency income shocks, the timing of many large parental transfers is significantly more predictable.

This paper joins the first and third literatures and analyzes the extent to which parents compensate for differences in their schooling investments in their children.[5] In this study, we assume that children's initial cognitive abilities are exogenous, and that their incomes are endogenous. Building on the first literature, our model predicts that children with higher abilities acquire more schooling. Building on the third literature, we study how parents use transfers at predictable times to mitigate earnings inequalities generated by unequal schooling investments.

We investigate this behavior using data from a survey of rural households that we carried out in Hebei, China. There are several reasons why our framework is appropriate for this setting. First, Chinese rural villages are primarily agricultural and patrilocal. Sons usually either live with their parents or remain in the same village after they marry. Parents and adult sons also often engage in farming and non-agricultural activity together, suggesting much closer economic interactions than observed in modern urban environments. In this setting, individual earnings and

[1] e.g., Behrman et al. (1994); Plug and Vijverberg (2003); Qian (2008).

[2] e.g., Altonji et al. (1997); Hochguertel and Ohlsson (2009); McGarry and Robert (1997, 1995); Nordblom and Ohlsson (2011); Wolff et al. (2007).

[3] e.g., Hochguertel and Ohlsson (2009); McGarry and Robert (1997, 1995); Wolff et al. (2007).

[4] Menchik (1980, 1988) argued that most parents divided their estates equally among the children. Tomes (1981) disagreed. Using a sample of wealthy US families, Wilhelm (1996) showed that most parents give equal bequests to their children. Among the children receiving unequal parental bequests, the differences are related to differences in their incomes.

[5] Our concerns are largely unrelated to the second literature which studies family insurance against exogenous high frequency (annual) income shocks.

consumption are also usually unobserved. Analyzing allocations using a total family wealth framework as we do here seems reasonable. Second, in our data, educational expenditures are the largest component of monetary parental investment. Bride-prices and land bequests are typically the largest transfers from parents to sons. Thus, our framework captures the most important inter vivos transfers in this setting.[6]

The first part of the paper uses a common behavioral framework to derive two alternative methods for estimating the effect of differences in schooling expenditure on parental transfers to their sons. The first method is the log-linear family fixed effect model. Although widely used in the literature, this empirical model has not been behaviorally derived. We show that a critical assumption for this model is that a child retains essentially all of his/her own labor earnings. This assumption is more likely to hold in modern industrial societies than the patriarchal rural society considered here, where co-habitation between parents and sons is very common. Our second method is a multiplicative fixed effects model, where family fixed effects are interacted with observable sibling characteristics in the transfer equation. This model does not make any assumption on the share of children's labor earnings appropriated by the parents.

To test these two models empirically, we focus on a parameter capturing the additional transfers a son receives from his parents when they invest one more yuan in schooling on his brother. We call this parameter the "marginal compensation coefficient." Most empirical work relies on data on partial lifetime transfers. Under the multiplicative model, we show that the marginal compensation coefficient should increase as the observable fraction of lifetime transfers increases; however, under the log-linear family fixed effect model, the coefficient should not change. This restriction on the log-linear model is new to the literature. The implementation of this test crucially depends on two features of our data. First, we observe multiple children within the same household so that we can control for the unobservable household-level heterogeneity. Second, we observe more than one significant transfer from parents to children.

Estimation of the log-linear household fixed effects regression of marital transfers on schooling expenditure delivers an estimated marginal compensation elasticity of 0.26. By adding post-marital land transfers to the bride-price, the value of the estimated compensation elasticity increases to 0.4. The conventional log-linear specification is strongly rejected by our data, and in the rest of the paper, we focus on estimating our multiplicative family fixed effects specification.

Theoretically, the marginal compensation coefficient is the product of three factors: the fraction of lifetime transfers observed by the researcher; the share of labor earnings of the son appropriated by the family; and the gross return, principal plus interest, to schooling expenditure. Our empirical results indicate that when a son receives 1 yuan less in schooling investment than his brother, he will obtain 0.47 yuan more in marital transfers and land division as partial compensation. Since marital transfers and land bequests are the largest transfers these parents will make to their children, the 0.47 point estimate suggests that parents partially, but not fully, compensate children for inequality in schooling investments. This also implies that lifetime intra-household consumption across siblings favors the child with more education. We reject Becker's unitary model where parents have equal concerns for their children's consumption.

[6] Parents regularly provide/transfer resources to children but most of these frequent inter vivos transfers are small.

We check the robustness of our results along several dimensions. Accounting for measurement error in schooling expenditure, or discounting of the timing of school expenditures and marriage transfers does not substantially change our results. Our Monte Carlo simulation results also show that the potential bias associated with the "incidental parameter problem" using multiplicative fixed effects is not significant.

Our main finding that large inter vivos transfers are partially compensatory for past unequal parental investments is consistent with work on the USA (see Dunn and Phillips 1997; McGarry and Schoeni 1995; and Hochguertel and Ohlsson 2009) and France (see Wolff et al. 2007). Using related but different "within-family" estimators, these authors find that inter vivos parental gifts partially compensate for unequal permanent incomes among siblings.[7]

Our paper is also related to a growing literature on intra-household resource allocation in rural China where nearly half of the population still lives. Qian (2008) found mixed support for the unitary model in her study of gender differences in parental investments. Wei and Zhang (2011) suggest that competition for brides, mainly driven by imbalances in the local sex ratio, may provide stronger incentives for parents to save and invest in their sons than their daughters. Such an effect is greater in rural China than its urban counterpart. A related paper on parental compensation which deals with a very different prior parental decision is by Li et al. (2010). They studied how parents compensated children who they were forced to choose for involuntary rustication during the Cultural Revolution (1966–1976). Using a dataset of monozygotic twins, they find that urban parents provided larger marital gifts to the twin who the parents chose to send down to the countryside relative to the twin who was able to remain at home.

This paper is organized as follows. Section 2 presents a model of parental investment, derives testable implications, and discusses identification issues. Section 3 introduces the data and descriptive results. Section 4 reports the estimation results. Section 5 addresses some related issues. And Section 6 concludes the paper.

2 An empirical framework

In this section, we sketch a simple model of the household to derive our estimating equations. Consider a household h with two children, denoted by i=1, 2, respectively. The parents have an initial wealth endowment of m_h. Child i has an initial cognitive ability level of a_{ih}. To generate family wealth, parents can invest in their children's schooling. Let the parents spend s_{ih} on child i's schooling. This expenditure generates $R(s_{ih})$ revenue from i's schooling for the family. In rural households, $R(.)$ is the contribution of the child to family income. If the child works outside the home, $R(.)$ would be labor earnings.

The total cost of s_{ih} to the family is $C(s_{ih},a_{ih},m_h)$ where schooling expenditure, s_{ih}, is determined by a child's initial cognitive ability, a_{ih}, and the initial wealth endowment of the family, m_h.

[7] Significant differences in environments and the fraction of lifetime transfers that are observed between our study and theirs make these comparisons less than perfect.

The budget constraint of the family is

$$c_h + c_{1h} + c_{2h} = m_h + \sum_{i=1}^{2} [R(s_{ih}) - C(s_{ih}, a_{ih})] \equiv w_h$$

where c_h represents the consumption of the parents; c_{ih} is child i's consumption; and w_h denotes total family wealth.[8]

For a large class of models, it can be shown that optimal consumption for each family member is proportional to the total family wealth[9]:

$$c^*_{1h} = k_{1h} w_h \tag{1}$$

$$c^*_{2h} = k_{2h} w_h \tag{2}$$

$$c^*_h = (1 - k_{1h} - k_{2h}) w_h \tag{3}$$

where k_{ih} is the share of family wealth allocated to child i, i=1, 2.[10] In the benchmark case of Becker's unitary model with equal parental concerns for both children, k_{1h}= k_{2h}=k_h. In general, we expect $k_{1h} \neq k_{2h}$. Differences in these shares may reflect parental preferences, bargaining power of siblings over family resources, and future exchange considerations between parents and children.

Each of these effects will be influenced by the labor earnings of the child. To capture this dependence, consider the linear projection of k_{ih} on $R(s_{ih})w_h^{-1}$:

$$k_{ih} = k'_{ih} + \mu \frac{R(s_{ih})}{w_h}; \quad i = 1, 2 \tag{4}$$

where $R(s_{ih})w_h^{-1}$ is the child's labor earnings as a share of total family wealth, and μ is the proportion of a child's earnings retained by the child.

By substituting Eq. (4) into Eqs. (1) or (2),

$$c^*_{ih} = k'_{ih} w_h + \mu R(s_{ih}) \tag{5}$$

Child i's consumption is the sum of two terms. The first term is proportional to total family wealth, w_h, while the second is the share of own labor earnings retained by the child.

Equation (5) provides an economic interpretation of μ, which measures the increase in child i's lifetime consumption due to a marginal increase in their lifetime earnings holding the total family wealth constant. Since the total family wealth is held constant here, a marginal increase in i's income implies a comparable reduction in other family members' total income. The same is true for increases in their consumption. In other words, μ is an intra-household redistributive parameter.[11]

[8] When a family has more than two children, consumption of the other children is subsumed in parental consumption.
[9] In Appendix A, we derive the optimal consumption under the unitary and collective model.
[10] The fact that optimal consumption is a fixed proportion of the total family wealth comes from the conventional homothetic assumption on the household objective function (e.g., Becker 1991). See Appendix A for details.
[11] It is similar in spirit to the inter-temporal labor supply elasticity which holds the marginal utility of wealth constant. In that context, lifetime wealth is held constant and the elasticity measures how current labor supply responds to an increase in the current wage relative to wages in all other periods.

Becker's benchmark unitary model where parents do not favor any particular child implies μ=0 and k'_{ih}=k_h. A unitary model where the parents favor the child with higher earnings will result in μ>0. For example, parents that are proud of the higher-earning son may decide to provide him a larger bride-price and level of consumption than his brother. In this case, μ>0, even though the parents have full control over the distribution of intra-household resources.

In most bargaining and strategic models of household behavior, μ>0. Thus, μ>0, by itself, cannot distinguish between the unitary and other models of intra-household consumption. In fact, a common test of the unitary model is to test whether the consumption of a household member is independent of that individual's contribution to household wealth, i.e., whether μ=0. For such a test to be valid, the researcher must add the assumption that the household utility function does not favor that individual, i.e., k'_{ih}=k_h.[12]

When μ=1, the son consumes:

$$c^*_{ih} = k'_{ih} w_{ih} + R(s_{ih}) \tag{6}$$

In this case, he fully expropriates the marginal increase of his contribution to family wealth. But as long as k'_{ih} is not equal to zero, family wealth still affects the child's consumption beyond his own earnings.

2.1 Transfers and schooling expenditures

A child i's lifetime consumption is unobserved, and we only observe a portion of total parental transfers to children. Given the child's consumption, Eq. (5), lifetime parental transfers, t_{ih}, are:

$$t_{ih} = c^*_{ih} - R(s_{ih}) = k_{ih} w_h - R(s_{ih}) \tag{7}$$

$$= k'_{ih} w_h - (1-\mu) R(s_{ih}) \tag{8}$$

The above equation says that the lifetime parental transfer is equal to the share of family wealth going to the child minus the share of child i's earnings, $1-\mu$, appropriated by the family. Researchers do not observe lifetime transfers between parents and their children. At best, they may observe several significant transfers. Let the researcher observe the fraction α of lifetime transfers. Equation (8) then becomes:

$$\alpha t_{ih} = \alpha k'_{ih} w_h - \alpha(1-\mu) R(s_{ih}) \tag{9}$$

where αt_{ih} is the amount of observed transfer. We now provide two ways to estimate Eq. (9).

[12] Lundberg et al. (1997) have a nice test of this nature. They show that the consumption basket of a household changes when the government changes the assignment of a family subsidy from the husband to the wife. In this case, the change in the distribution of the contribution to family wealth by household members has changed in a lump sum manner but the household utility function should remain unchanged under the unitary model.

2.2 Log-linear fixed effects models

Most existing studies consider a log-linear family fixed effects version of the transfer equation:

$$\ln \alpha t_{ih} = \ln\omega_h + X_{ih}\rho - \beta \ln s_{ih} + \varepsilon_{ih} \tag{10}$$

where ω_h is a measure of household wealth, captured by the household-fixed effect. In addition, there are observed fixed differences across children, X_{ih}, such as birth order. These observable differences may reflect k_{ih}. β is the elasticity of transfers with respect to expenditure on child i's schooling, s_{ih}. Although versions of the above model are commonly used, it has never been behaviorally justified. We provide a justification here.

Taking logs of the transfer determination Eq. (9), we have:

$$\ln(\alpha t_{ih}) = \ln\left[\alpha k'_{ih} w_h - \alpha(1-\mu)R(s_{ih})\right]$$

Applying a first-order Taylor series expansion of μ around 1 to the right hand side, we have:

$$\begin{aligned}\ln\left[\alpha k'_{ih} w_h - \alpha(1-\mu)R(s_{ih})\right] &\approx \ln\left[\alpha k'_{ih} w_h - \alpha(1-\mu)R(s_{ih})\right]\big|_{\mu=1} \\ &\quad + \frac{\alpha R(s_{ih})}{\alpha k'_{ih} w_h - \alpha(1-\mu)R(s_{ih})}\Big|_{\mu=1}\cdot(\mu-1) \\ &= \ln\left(\alpha k'_{ih} w_h\right) - (1-\mu)\frac{R(s_{ih})}{k'_{ih} w_h}\end{aligned}$$

Approximating $R(s_{ih})(k'_{ih}w_h)^{-1} = \gamma \ln s_{ih}$, we get:

$$\ln \alpha t_{ih} = \ln\alpha + \ln w_h + \ln k'_{ih} - \beta \ln s_{ih} + \varepsilon_{ih} \tag{11}$$

$$\beta = \gamma(1-\mu) \tag{12}$$

Our derivation of the log-linear fixed effects model is based on two strong assumptions: First, $\mu \approx 1$, which implies that a son retains all his labor earnings. In the context of rural China where sons often work on the family farm and in related family businesses, this assumption may be unreasonable. Interestingly, it may be more appropriate in modern urban economies where sons work separately from their fathers. And second, $R(s_{ih})(k'_{ih}w_h)^{-1} = \gamma \ln s_{ih}$. When $\mu \approx 1$, by Eq. (5), $R(s_{ih})(k'_{ih}w_h)^{-1}$ is the ratio between the son's consumption out of his labor earnings and that related to his share of family wealth. We expect this ratio to grow with his earnings, however, assuming that it grows proportionately to the log of schooling expenditure is only one of several alternative possibilities.

Equation (11) also implies that the compensation elasticity with respect to schooling expenditure, $\beta = \gamma(1-\mu)$ is independent of α. That is, adding more observed transfers to the dependent variable, or increasing α does not affect our estimate of the lifetime compensation elasticity. This is a testable restriction of the log-linear fixed effects model that we will examine later in this paper.

2.3 Multiplicative fixed effects model

Starting again from the transfer determination, Eq. (9), for empirical tractability, we let $R(s_{ih})=r_0+rs_{ih}+\varepsilon'_{ih}$, where r, the return to schooling, is assumed to be homogenous across siblings.[13] We expect $r>1$ because this is the gross return, principal plus interest, to schooling expenditure.

We then have:

$$\alpha t_{ih} = \alpha k'_{ih} w_h - \alpha(1-\mu) r s_{ih} + \varepsilon_{ih} \tag{13}$$

Given a random sample of households $h=1, .., H$, consider the regression

$$\alpha t_{ih} = \kappa_{ih} F_h - \beta s_{ih} + \varepsilon_{ih} \tag{14}$$

where $\{F_h\}_{h=1}^{H}$ are parameters capturing household wealth, κ_{ih} is the proportion of household wealth allocated to each son for consumption purpose, β is the marginal compensation coefficient, and ε_{ih} is the IID measurement error in transfers. β estimates the amount by which transfers, αt_{ih}, are reduced for each additional yuan of schooling expenditure. The main difference of this regression with the log-linear form conventionally used in the literature is that household fixed effects enter the regression in a multiplicative form.

Comparing Eqs. (14) with (13), β estimates $\alpha(1-\mu)r$. Since $\alpha(1-\mu)r$ is positive, we expect $\beta>0$. As discussed above, $r>1$, $\alpha<1$, and $(1-\mu)\leq 1$. As a result, $\alpha(1-\mu)r$ is of indeterminate magnitude. Note also that β is increasing in α. Thus, as the researcher observes more lifetime transfers, the estimated marginal compensation coefficient should increase.

In summary, this section presented two empirical models of inter vivos transfers to sons. Both models make different functional form assumptions about $R(s_{ih})$. The log-linear model has an additional assumption that μ is close to 1. The models also differ in their predictions on how partially observed lifetime transfers to one son change in response to an increase in schooling expenditure to the other son. Under the log-linear model, the compensation elasticity is independent of α. Under the multiplicative fixed-effects model, the marginal compensation coefficient is increasing in α. We estimate both models and test the restrictions implied by each model.

Both models are family fixed-effect models. Thus, our methodology has nothing to say about compensation for unmarried sons. As long as the family has at least two married sons, we can use this family to consistently estimate our model independent of how many unmarried sons it also has.

3 Data

The data used in this study come from the "Survey on Family and Marriage Dynamics in Hebei Province," which was carried out in rural Hebei in the

[13] In Section 4.4 we relax the homogeneity assumption and discuss how it will affect our results.

summer of 2005 by the authors and their Chinese colleagues. The general purpose of this survey was to investigate how three decades of state-initiated economic, social, and political change have affected marriage institutions and families in rural China. Rural Hebei province is culturally, economically, and socially representative of North China. Altogether, 600 households from 30 villages, in 15 townships in 3 counties were surveyed.[14] Figure 1 locates each of the counties in Hebei.

Each household was required to have at least one married child in order to be included in the sample. Our respondents were parents in the household between the ages of 50 to 69. We excluded parents older than 70 because of concerns about recall. For each child, information on significant events over his/her life was collected from the parents, including (1) education, (2) pre-marital work experience, (3) engagement and marriage, (4) fertility and post-marriage intra-family arrangements, and (5) pre-mortem household division where applicable. To minimize the burden of the interview, for households with more than three married children, three of them were selected for detailed enumeration.[15]

3.1 Sons' sample

In this paper, we focus on the multiple-son sample in our dataset, i.e., households with more than one married son for whom we have complete information.[16] This is based on several considerations. First, when both sons receive positive transfers, transfer decisions reflect interior solutions to the parents' optimization problem. In our data, nearly all sons receive strictly positive marital transfers; however, the same is not true for daughters for whom about 20 % do not receive a dowry.[17] Second, because of the nature of patrilocal society and village land rules, only sons are entitled to receive household land as part of household divisions. As a result, our test of the log-linear model, which requires information on more than one transfer, is only feasible for the sons' sample. Finally, the magnitude of the bride-price and sons' eligibility to receive land transfers suggests that our measure

[14] The three counties, Fengrun, Zhaoxian, and Chicheng, were selected after extensive analysis of county and township level economic and demographic information from the 1980s to the 1990s. Fengrun is the richest of the three in terms of per capita GDP, and Chicheng is the poorest. Within each county, townships were ranked on the basis of incomes, and then one was randomly selected from each quintile. Two villages were then randomly selected in each township: one from the upper half of the income distribution, and one from the lower half. Finally, from each village, 20 households satisfying age requirements for the household head and his spouse were randomly selected.

[15] There are, in total, 103 households (17 %) having more than three married children. The selection criteria for picking children in these households were: (1) Choose the first and the most recently married child; (2) If the two selected children are of the same sex, choose a third of the opposite sex (43 households); and (3) If the two selected children are of the opposite sex, we randomly selected a third child (60 households).

[16] In the multiple-son sample used in our study, the minimum age is 22. Since we focus on the married sons, we are ignoring those adult sons (older than 18) in the sampled family but remained unmarried. However, they only account for 2 % of all adult sons in the sampled family.

[17] For these cases, implications of the model's corner solution need to be derived, which is the basis for future work.

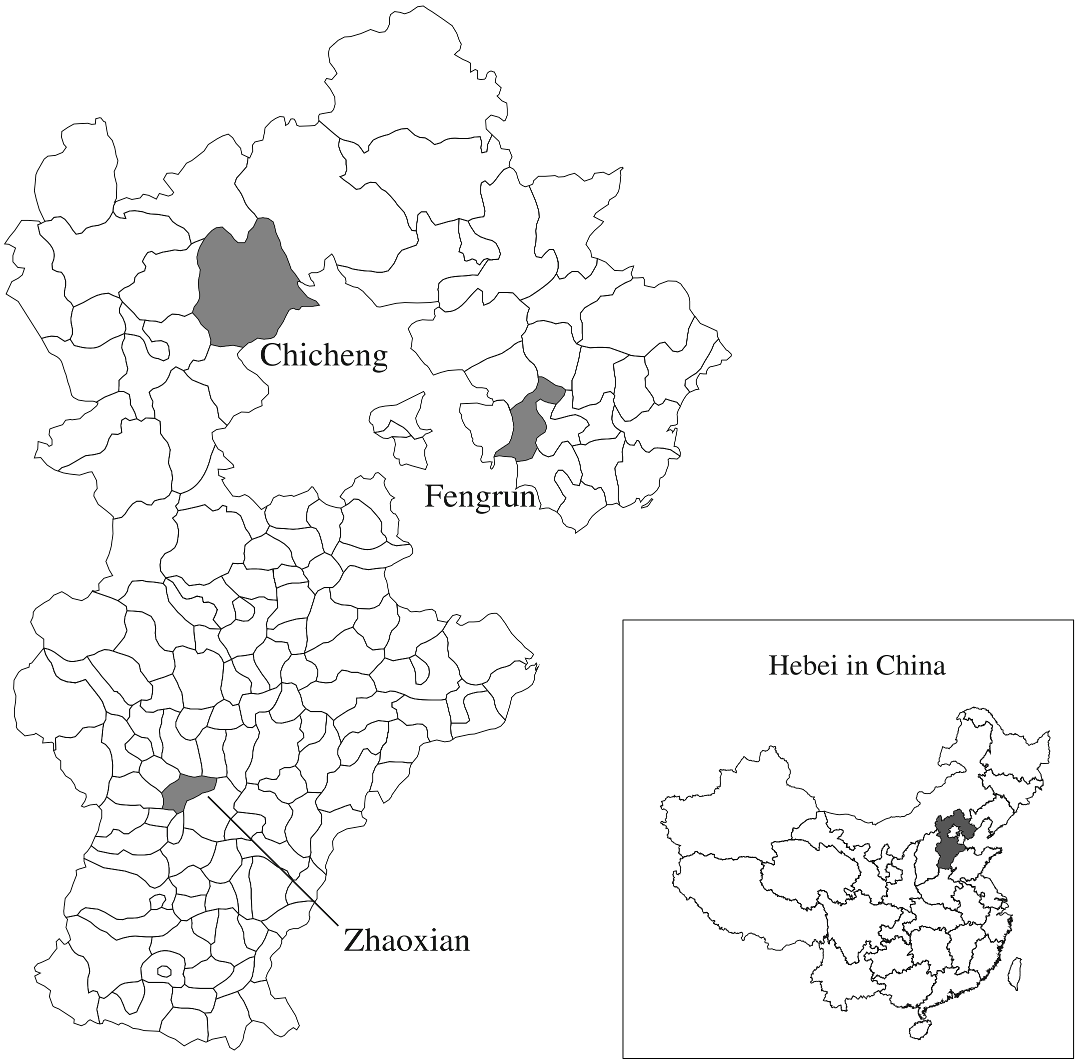

Fig. 1 A map of Hebei locating the three counties that were surveyed in this study. The three surveyed counties are *shaded* in. The large *white area* in the middle of the province is the municipalities of Beijing and Tianjin, both of which have provincial administrative status

of total inter vivos transfers is a better measure of lifetime transfer for sons than for daughters.

In total, the multiple-son sample includes 140 households with 293 married sons. Their basic summary statistics are reported in Table 1.

3.2 Education expenditure

In rural Hebei, parents shoulder almost all of the educational costs of their children. Self-financing by children is negligible. For each child in our dataset, we collected data on educational expenditure since middle school (grade 7), including tuition and fees, books, and room and board, all deflated to 1980 price levels. We do not have expenditure data for elementary school because most parents could not accurately recall such information. Instead, we impute the spending on elementary school for each child. A study by Liu et al. (2006) shows that the total cost of elementary education is about half the cost of middle school education in rural China. We regress-observed middle school expenses on gender, cohort, and village dummies,

Table 1 Variable definitions and summary statistics

Variable name	Unit	Definition or notes	Observations	Mean	Standard deviations	Standard deviations within households	Number of households with intra-households variations
Key variables							
Educational expenditure	Yuan[a]	Total educational expenditure	293	1,478	2,661	1,829	127
Marriage transfers	Yuan	Total monetary value of bride-price, including house, items, and cash	293	5,064	5,349	2,655	127
Land division	Yuan	Value of the land division[b]	186	2,221	1,989	987	83
Sons' attributes							
Age		Age	293	33.73	5.66	2.94	141
Age at marriage		Age at marriage	293	23.49	2.63	1.72	119
Height	cm	Height	293	170.37	5.86	2.78	110
Agricultural experience	Year	Agricultural experience before marriage	293	3.53	4.53	2.27	82
Non-agricultural experience	Year	Non-agricultural experience before marriage	293	3.17	3.36	2.00	101
Years of schooling	Year	Years of schooling.	293	8.40	2.92	1.64	83
Dummy variables used in the specifications							
D_Height		Indicator of the taller son[c]	293	0.39	0.49	0.44	110
D_Ag_Experience		Indicator of the son with more years of experience in agriculture	293	0.30	0.46	0.38	82
D_Nonag_Experience		Indicator of the son with more years of experience in non-agriculture	293	0.35	0.48	0.42	101
D_Live_with_Parents		Indicator of living with parents after marriage	293	0.36	0.48	0.30	52
D_1st_Son		Indicator of the first born son	293	0.47	0.50	0.49	138

Table 1 (continued)

Variable name	Unit	Definition or notes	Observations	Mean	Standard deviations	Standard deviations within households	Number of households with intra-households variations
Household attributes							
Father's years of schooling	Year	Father's years of schooling	140	5.20	3.16		
Number of sons		Total number of sons in the household	140	2.43	0.76		
Number of daughters		Total number of daughters in the household	140	0.82	0.91		
Value of houses	Yuan	Total value of houses ever built by the parents	140	11,151	11,906		
Value of agricultural equipment	Yuan	Total value of agricultural equipment ever possessed by the parents	140	2,100	4,849		
Fengrun		Indicator of residents in Fengrun county	140	0.33	0.47		
Zhaoxian		Indicator of residents in Zhaoxian county	140	0.27	0.45		
Chicheng		Indicator of residents in Chicheng county	140	0.39	0.49		

[a] All monetary values are deflated to the 1980 price level

[b] Construction of land value measurement. Hebei Statistics Yearbook provides the net income from cultivation per mu in each year. We assume that half this income is the net return to human capital. Given the area of the land received by each son, the value of the land is defined as the sum of the total return to human capital over the 10 years since the land was given by the parents

[c] Based on a comparison among the selected sibling in the household. "D_Ag_Experience" and "D_Nonag_Experience" are defined in a similar manner

and halved the predicted value before using it as imputed elementary schooling expenditure.[18]

Figures 2 and 3 illustrate the dynamics of real educational expenditures in our sample. Following economic reform, total educational expenditure on sons increased significantly along with household annual income. These expenditures rose especially fast early on, but the increase slowed beginning in the late 1980s. Total household expenditure on a son's education amounts to about four-fifth of the annual household income. Increases in educational expenditures arose mainly from an increase in years of schooling rather than increases in annual costs (see Fig. 3).[19]

3.3 Marital transfers

A quantitatively significant inter vivos transfer in our dataset is marital transfers. Currently, a large part of the marital transfers in rural Hebei is given by parents to their marrying children, and not to their in-laws. It is an intergenerational transfer rather than an interfamilial transfer. At the time of the marriage, the groom's and the bride's families provide furniture, major home appliances, farm equipment, and sometimes cash payments to the newly-wed couple.[20] The groom's family usually spends more because traditionally, they are responsible for building a new house for the newlyweds. In our dataset, for each marriage, we have a complete inventory of marital transfers along with their monetary value. The average of these transfers was 5,085 yuan.

Parents save for years in order to finance these expenditures. Escalating marital expenses in Chinese villages during the past two decades have been well documented in the literature (e.g., Siu (1993); Min and Eades (1995); Zhang (2000); Wei and Zhang (2011)). Our dataset confirms high and rising marital transfers, especially in bride-prices, since the economic reform (see Fig. 2.) In most of the period we examined, the bride-price is roughly two and a half times the annual household income. With reported rural household savings running about 30 % of reported household income,[21] parents must save for 7 to 8 years to accumulate the bride-price. By contrast, the dowry is typically equal to about half of the annual household income.

3.4 Land division

Some families in our sample made household property divisions before the parents were deceased. These agreements were typically verbal, with only a fifth relying on

[18] The mean of this imputed variable is around 300 yuan, which represents approximately 20 % of average educational expenditure on sons.

[19] We also examine educational expenditures on daughters over the same period. In the 1980s, spending on girls' education was below that on boys, but by the mid-1990s, investments in children's human capital are about the same between genders.

[20] The cash transfer paid by the groom's family, which accounts for about 20 % of the bride-price, is usually given directly to the bride's family. However, once the bride's parents receive this cash payment, they can either keep it or use it to purchase items for the dowry. The dowry financed by the cash component of the bride-price is often referred to as "indirect dowry" in the sociology literature (see Goody (1973)). According to our data, on average, nearly all of the cash in the bride-price became indirect dowry. This suggests that the cash component in the bride-price should be included in the marital transfers enjoyed by the grooms, even though they are initially given to the bride's parents as inter-familial transfers.

[21] National Bureau of Statistics, China Statistical Yearbook (2007)

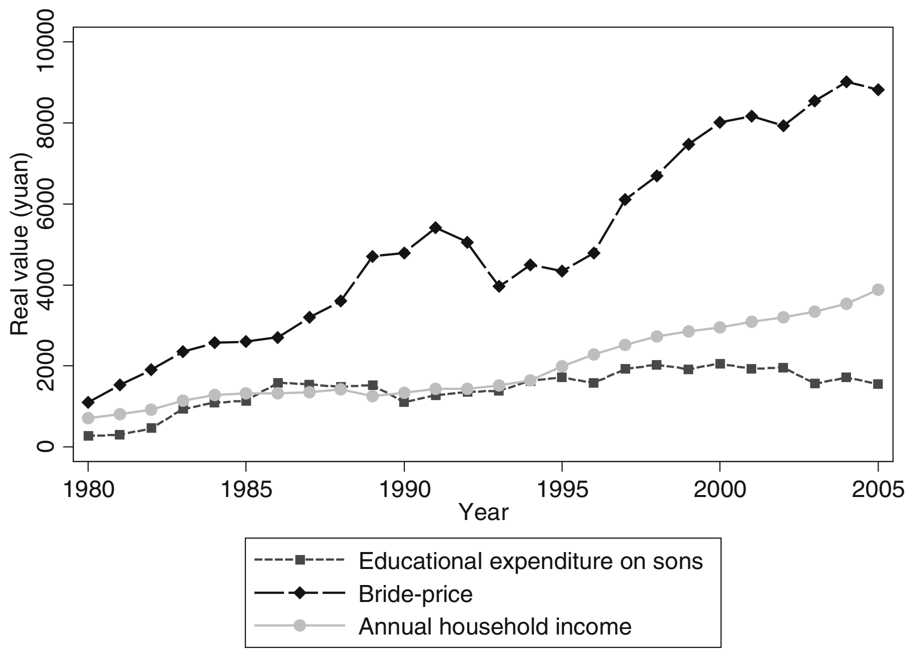

Fig. 2 Change in marital transfers (bride price) and educational expenditure in this study sample since the economic reform. Annual household income data is obtained from the Hebei Bureau of Statistics (http://www.hetj.gov.cn/); marital transfers and educational expenditure comes from the Survey on Family and Marriage Dynamics in Hebei province conducted by the authors

formal written contracts. Daughters in rural China are not entitled to any forms of (pre-mortem) bequests. The pre-mortem bequests include housing, land, and other producer durable goods, but the housing transfers are already included in the bride-price. For land divisions, we know the amount of the land and the time of transfer. For purposes of valuing the inheritance associated with land, a few institutional details regarding property rights are required.

Fig. 3 Dynamics of real educational expenditures in this study's sample

Land in rural China is not privately owned. Rather, ownership rights reside with the village (collective), and households are allocated usufruct rights. Since the introduction of economic reform in the late 1970s, these rights have largely been allocated to households on a per capita basis, with the length of tenure governed in principal by a series of land laws. For example, the first national land law extended tenure for 15 years. The land laws have not always been respected, and land has often been reallocated among households by villages before the expiration of the land law, however in Hebei province, use rights have been very secure.

As part of the household division of property, sons are effectively receiving usufruct rights to the land that has been allocated to the household by the village. At a minimum, these rights are secure until the next land law, but in practice, the claims run much longer. Conservatively, we calculate the value of the land as the total real return to the land from cultivation over the first 10 years after the land is given to the son.[22]

At the time of our survey, 63 % of our households had already divided the land, with a mean value of 2,221 yuan (see Table 1). Only about 10 % of the sampled households report dividing items other than land and housing, with a mean value of 952 yuan. Based on these statistics, land is the only other important transfer besides marital transfers that is captured by our data.[23] [24]

3.5 Descriptive statistics

Table 1 presents descriptive information on the characteristics of these households and their children. To illustrate the relative importance of the between-sibling variation in our dataset, we calculate for individual-level variables the within-household standard deviations in column 4. The results are generally half of the overall standard deviations (column 3 of Table 1). Column 5 reports the number of households that have within-household variation in children's characteristics.

Sons examined in this paper were born between the 1950s and the 1990s, with most marriages occurring between 1980 and 2000. On average, they had slightly more than 8 years of education, with a mean age at marriage of 23. Before marriage, sons typically had slightly more work experience in agriculture than non-agricultural activities.

The education of the father is a potentially important determinant of parental wealth, and on average, was slightly more than 5 years. In some regressions, we use the sum of the total real value of housing and fixed agricultural assets as a proxy for family wealth. These estimates are constructed on the basis of a

[22] We assume that the share of land in value-added is 50 %. Because of discounting, increasing the number of years does not significantly change the value of the inheritance.

[23] In our survey, we asked parents about any other transfers larger than 100 yuan given to their children after they marry. Only 43 sons (16 % of the total sample of sons) received other post-marital transfers, with an average of 1,700 yuan. This value is only about a third of the mean bride-price these sons received. The evidence suggests that large post-marital transfers are not a common practice in rural Chinese villages.

[24] We also estimate our model adding these post-marital transfers and the value of assets other than land received as part of household division to our new measure of inter vivos transfers. The results are not significantly changed from those obtained only by using land and the bride-price.

Table 2 Within households comparison of sons' characteristics

Characteristics	Unit	Full multiple-son sample			Sample with land divisions		
		More educational expenditure	Less educational expenditure	*P* value	More educational expenditure	Less educational expenditure	*P* value
		(1)	(2)	(3)	(4)	(5)	(6)
Years of Schooling	Year	9.27	7.53	0.00	8.92	7.16	0.00
Height	cm	170.16	170.57	0.54	169.89	170.57	0.44
Agricultural experience	Year	2.82	4.24	0.01	3.13	5.15	0.00
Non-agricultural experience	Year	3.22	3.12	0.81	2.78	2.99	0.67
Live with parents		0.33	0.39	0.24	0.37	0.44	0.29
1st Son		0.46	0.47	0.86	0.49	0.45	0.57
Observations		147	146		93	93	

The table reports a within-household comparison of sons receiving more in educational expenditure (columns 1 and 4) with those receiving less (columns 2 and 5). Columns 1 to 3 use full multiple-son sample (293 sons in 140 households). Columns 4 to 6 use multiple-son sample with land divisions (186 sons in 89 households). Columns 3 and 6 report the *P* value in testing the null: "mean value of the two types of sons are not significantly different"

complete inventory of the investments in these assets made by the parents since they were married.

In Table 2, we compare the mean characteristics within the household of sons receiving higher educational expenditure with those receiving less. Columns 1–3 are based on the full multiple-son sample, while columns 4–6 are based on the sample with land divisions. In general, the results are very similar between the two samples. Not surprisingly, sons on whom parents spent less on education also have fewer years of schooling. Evidence also suggests that this reduction is schooling is associated with greater involvement in agriculture.

4 Empirical results

4.1 Log-linear fixed effects models

Our starting point is the log-linear regression model of parental transfers, Eq. (11), which is widely used in the literature (e.g., Wolff et al. 2007; Hochguertel and Ohlsson 2009[25]). Recall from Section 2.2 that if the log-linear transfer model is correct, β is independent of the fraction of lifetime transfers that is used as the dependent variable. Formally, suppose that we observe two

[25] Some of these papers use the inverse hyperbolic sine (IHS), $\sinh^{-1}(y_{ih})=\log(y_{ih}+(y_{ih}^2+1)^{1/2})$ transformation because it is similar to the $\ln(y_{ih})$ transformation and it admits zero transfers as a value.

inter vivos transfers t_1 and t_2, and denote β_{t_g} , where g=1,2 is the coefficient on schooling investment. When estimating Eq. (14) using transfer t_g as the dependent variable, support for the log-linear model is provided by a rejection of the null hypothesis

$$H_o : \frac{\beta_{t_3}}{\beta_{t_g}} > 1, g = 1, 2 \quad (15)$$

where $t_3=t_1+t_2$.

Panel A of Table 3 shows the coefficients on schooling expenditure ($-\beta_{t_1}$) in household fixed-effects regressions where the dependent variable is the log of marital transfers. In each of the columns, we control for different children characteristics. They include: (1) the son is taller than his sibling; (2) has more years of agricultural experience at the time of his marriage; (3) has more years of non-agricultural experience; (4) lives with the parents after marriage; and (5) is the eldest male. The estimated marginal compensation elasticity ranges from 0.20 to 0.33. Evaluated at the mean values of the regression sample, these estimates translate into a range of marginal compensation coefficients between 0.69 and 1.13. These estimates suggest that without any discounting, there is nearly full or full compensation. With modest discounting, these are less than full marginal compensation.

Panel B reports coefficients ($-\beta_{t_3}$) for the same specifications as panel A, but with the log of the sum of marital transfer and land division as the dependent variable. The estimated marginal compensation elasticity ranges from 0.40 to 0.58. Evaluating these coefficients once again at the mean values of the regression sample, they imply a marginal compensation coefficient in the range of 2.29 to 3.33. Without discounting, these estimates suggest more than full marginal compensation, contrary to the estimates on panel A. With modest discounting, the same holds true. Our log-linear results differ sharply from those reported in the literature, albeit in different contexts.[26]

Compared with panel A, estimates of the compensation elasticity β in panel B are almost double.[27] Given that β_{t_1} and β_{t_3} are estimated using different dependent variables, we rely on bootstrapping to test hypothesis Eq. (15). We first construct 100 bootstrapped samples (drawn with replacement) from households for whom we observe marital transfers (t_1) and also land as the main pre-mortem bequests to sons (t_2). We next estimate β_{t_1} and β_{t_3} for each of these samples, and then calculate the likelihood of β_{t_3}/β_{t_1} >1. Results in panel C suggest that we cannot reject hypothesis Eq. (15) at any conventional statistical levels. Since β is supposed to be independent of α, a significant increase of the estimated coefficient when α is increased is difficult to motivate under the log-linear model. Our explanation is that Eq. (10) is not a reasonable approximation of parental transfer behavior.

[26] See Dunn and Phillips (1997), McGarry and Robert (1997), and Hochguertel and Ohlsson (2009), using contemporary US data; and Wolff et al. (2007) using contemporary immigrant data in France. These papers use related log-linear but different "within-family" estimators.

[27] There are differences in the sample size between panels A and B. Therefore, we replicate panel A using the same restricted sample as panel B. We find that most of the coefficients are very similar, and only the coefficient on height, which becomes smaller in magnitude and insignificant, is sensitive to the use of the more restricted sample.

Table 3 Marital transfers regression with additive fixed effects (log-on-log form)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Panel A. Log(bride-price) as the dependent variable							
Log(edu_exp)	−0.31	−0.31	−0.26	−0.33	−0.22	−0.25	−0.20
$(-\beta_{t_1})$	[0.12]**	[0.12]**	[0.13]**	[0.12]***	[0.12]*	[0.12]**	[0.13]
D_Height		−0.12 [0.16]					−0.11 [0.15]
D_Ag_Experience			0.28 [0.21]				0.27 [0.22]
D_Nonag_Experience				−0.20 [0.17]			−0.14 [0.17]
D_Live_with_Parents					0.56 [0.29]*		0.62 [0.25]**
D_1st_Son						0.18 [0.14]	0.25 [0.14]*
R^2	0.67	0.67	0.67	0.67	0.68	0.68	0.69
Panel B. Log(bride-price + land) as the dependent variable							
Log(edu_exp)	−0.57	−0.58	−0.46	−0.58	−0.49	−0.57	−0.40
$(-\beta_{t_3})$	[0.13]***	[0.13]***	[0.14]***	[0.14]***	[0.12]***	[0.13]***	[0.14]***
D_Height		−0.17 [0.20]					−0.14 [0.18]
D_Ag_Experience			0.49 [0.22]**				0.47 [0.22]**
D_Nonag_Experience				−0.29 [0.22]			−0.14 [0.18]
D_Live_with_Parents					0.68 [0.27]**		0.66 [0.27]**

Table 3 (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
D_1st_Son						0.05	0.16
						[0.15]	[0.15]
R^2	0.60	0.60	0.62	0.61	0.63	0.63	0.66
Panel C. Bootstrap results							
Mean (β_{t_3}/β_{t_1})	1.28	1.45	1.57	1.29	1.50	1.28	2.10
Prob (β_{t_3}/β_{t_1} >1)	0.99	0.99	0.98	0.99	0.99	0.99	0.98

The table tests if $-\beta$ is invariant to the fraction of lifetime transfers that is used as the dependent variable. The dependent variable in panel A is the logarithm of the bride-price; in panel B, it is the logarithm of bride-price plus the value of land division

Panel A reports results using the full multiple-son sample (293 sons in 140 households)

Panel B reports results using the multiple-son sample with land divisions (186 sons in 89 households)

Panel C reports the mean of β_{t_3}/β_{t_1} and the probability of β_{t_3}/β_{t_1} >1 among 100 bootstrapped samples. Each sample draws 89 households with replacement using the original son sample with household divisions. Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

Our rejection of the log-linear model is not because we do not observe all the lifetime transfers which the sons receive. Rather, the log-linear specification has a well-founded behavioral interpretation when μ is close to 1. When μ is significantly less than 1, the log-linear specification does not have a well-founded behavioral interpretation independent of the value of α.

4.2 Main multiplicative fixed effects specification

By assumption, $k_{ih}^{'}$ is orthogonal to $R(s_{ih})w_h^{-1}$ in Eq. (4), however, it may not be orthogonal to s_{ih}. We add observed characteristics of son i in Eq. (14) to absorb the residual correlation between $k_{ih}^{'}$ and s_{ih}. Empirically, we allow κ_{ih} to depend on a set of binary indicators as described in Section 4.1. Estimation of κ_{ih} requires normalization. To ease exposition, let son i in household h be characterized by two indicator variables $D_{ih}=(D_{ih}^a,D_{ih}^b)\in\{0,1\}\times\{1,0\}$. Let the sons with $(D_{ih}^a,D_{ih}^b)=(0,0)$ be the base group denoted by κ_0. Not all households have a son with $\kappa_{ih}=\kappa_0$. For any type of son i,

$$\begin{aligned}\kappa_{ih}F_h &= \frac{\kappa_{ih}}{\kappa_0}\cdot\kappa_0 F_h \\ &= \left(1+\delta_a D_{ih}^a+\delta_b D_{ih}^b\right)\cdot\omega_h\end{aligned} \tag{16}$$

where the household fixed effects ω_h identifies $k_0 w_h$, which is the consumption level enjoyed by a child with hypothetical characteristics κ_0 in household h. δ_a or δ_b is the increase in k_{ih} due to the presence of characteristic a or b for child i relative to the hypothetical child 0 within household h. $\mu>0$ allows k_{ih} to also depend on s_{ih}.

Combining Eqs. (14) and (16), our empirical transfer specification is

$$\alpha t_{ih} = (1+D_{ih}\delta)\omega_h - \beta s_{ih} + \varepsilon_{ih} \tag{17}$$

The parameters $\left(\delta,\{\omega_h\}_{h=1}^{H},\beta\right)$ are estimated through non-linear least squares. Appendices B and C discuss the estimation algorithm and econometric issues including the incidental parameters problem.

Table 4 contains the main results of the multiplicative fixed effects model. Panel A reports regressions with the bride-price as the dependent variable. In all tables, the estimated coefficient on school expenditure is $-\beta$. We start with two simple level-on-level regressions which allow us to highlight the significance of having multiple sons in each household in order to control for unobserved household heterogeneities. In column (1), the point estimate of β using ordinary least squares (OLS) without any other covariates is 0.16 with a standard error of 0.07. Since marital transfers and schooling expenditure are positively related with household income, the estimate of β is biased downward without controls for household income. Column (2) adds household fixed effects to the regression while maintaining $k_{1h}=k_{2h}$, i.e., the equal concerns assumption. This yields an estimate of β of 0.34 with a standard error of 0.17. Controlling for household fixed effects, the estimated magnitude of β doubles

Table 4 Multiplicative specification

	OLS	Households fixed effects	Multiplicative fixed effects					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Bride-price as dependent variable								
Edu_exp	−0.16	−0.34	−0.35	−0.31	−0.33	−0.23	−0.33	−0.22
$(-\beta_{t_1})$	[0.07]**	[0.17]*	[0.13]***	[0.12]**	[0.12]***	[0.1]**	[0.12]***	[0.1]**
δ								
D_Height			0.24 [0.09]***					0.31 [0.09]***
D_Ag_Experience				0.16 [0.13]				0.00 [0.11]
D_Nonag_Experience					0.04 [0.09]			0.05 [0.1]
D_Live_with_Parents						0.56 [0.22]***		0.60 [0.16]***
D_1st_Son							−0.22 [0.1]**	−0.23 [0.09]**
R^2	0.01	0.75	0.77	0.76	0.76	0.78	0.78	0.82
F test $(P)^{\dagger}$								0.00
Panel B. Bride-price + land division as dependent variable								
Edu_exp	−0.42	−0.67	−0.67	−0.64	−0.66	−0.56	−0.63	−0.47
$(-\beta_{t_3})$	[0.07]***	[0.14]***	[0.14]***	[0.14]***	[0.14]***	[0.14]***	[0.14]***	[0.15]***
δ								
D_Height			0.01 [0.06]					0.03 [0.07]

Table 4 (continued)

	OLS	Households fixed effects	Multiplicative fixed effects					
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
D_Ag_Experience				0.06 [0.1]				0.05 [0.11]
D_Nonag_Experience					0.04 [0.07]			0.05 [0.09]
D_Live_with_Parents						0.20 [0.11]*		0.21 [−0.19]
D_1st_Son							−0.26 [0.08]***	−0.19 [0.08]**
R^2	0.03	0.81	0.81	0.81	0.81	0.82	0.83	0.84
F test $(P)^{\dagger}$								0.01
Panel C. Bootstrap results								
Mean $(\beta_{t_3}/\beta_{t_1})$			1.89	1.80	1.88	2.18	2.20	3.26
Prob $(\beta_{t_3}/\beta_{t_1} > 1)$			1.00	1.00	1.00	1.00	1.00	0.99

Panel A reports results using the full multiple-son sample (293 sons in 140 households)

Panel B reports results using the multiple-son sample with land divisions (186 sons in 89 households)

Panel C reports the mean of β_{t_3}/β_{t_1} and the probability of $\beta_{t_3}/\beta_{t_1} > 1$ among 100 bootstrapped samples. Each sample draws 89 households with replacement using the original son sample with household divisions

Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

† P value of the F test for the joint significance of k_{ih}

in the direction predicted by the theory. Fitness of the model also significantly increases.[28]

Next, we estimate Eq. (14), the multiplicative fixed effects version of our model. Columns (3) to (8) in panel A of Table 4 report the estimation results using various children's attributes as controls for k_{ih}. Throughout different specifications, the point estimate for β generally remains in the vicinity of 0.3. Coefficients on each covariate (δ) can be interpreted as the difference in consumption between sons with such an attribute and a "hypothetical" son in the same household, for whom all categorical variables take on zero values. We use column (8) with full controls as our preferred specification. Here, compared with the "hypothetical son" in the household, taller sons enjoy one third more in total consumption,[29] either because parents put greater weights on children with stronger physical endowments, or because these children tend to have more bargaining power on household consumption allocation. Sons who live with their parents after marriage obtain nearly 61 % more consumption than those who do not, which is consistent with the exchange motives of the inter vivos transfers. In contrast, the eldest sons tend to receive one fifth less than their younger siblings, which at first sight, is inconsistent with popular wisdom about first-son bias in rural China. In Section 5.1, we show that this inconsistency is partially resolved with appropriate discounting. We do not find any evidence that sons with more agricultural or non-agricultural experience enjoy advantages in total consumption. All together, these covariates are jointly significant and adding them decreases our compensation coefficient by one-third.

Estimates of $k_0 w_h$, consumption for the hypothetical baseline child in household h, are reported in Table 5. The results are very intuitive. Panel A shows the summary statistics of $k_0 w_h$ estimated under alternative specifications in Table 4. Figure 4 display its distribution.[30] Both the figure and the Kolmogorov-Smirnov statistics show that the distributions of $k_0 w_h$ are not significantly different from log normal. In panel B, we examine the determinants of $k_0 w_h$, by regressing it against household-level characteristics. The estimates are consistent across columns. The coefficients on the number of sons and daughters in the household are significantly negative, suggesting that the hypothetical son enjoys less consumption if there are more children in the family. Coefficients on the total value of houses ever built by parents and the total value of agricultural equipment are significantly positive, indicating that the level of consumption is higher for children from wealthier families. Moreover, we also find weak evidence suggesting that children from families with better educated household heads, or living in the richer county (Fengrun and Zhaoxian) enjoy higher consumption.

4.3 Adding land division

To see how an increase in α affects the estimation of β in the multiplicative fixed effect model, we add the value of land divisions received by sons to the dependent variable.

[28] R-squared increases from 0.01 in column (1) to 0.75 in column (2)

[29] Aside from genetic factors, variation in heights may be due to time-varying exogenous environmental factors such as the Great Leap Famine, agricultural reform, etc., as well as family-level shocks.

[30] $k_0 w_h$ in this figure are estimated from specification (8) in panel A, Table 4.

Table 5 Properties of k_0w_h

	Specification in Table 4 used to estimate k_0w_h					
	(3)	(4)	(5)	(6)	(7)	(8)
Panel A. Summary of k_0w_h from specifications in Table 4						
Observations	140	140	140	140	140	140
Mean	5,052	5,232	5,401	6,087	4,124	4,260
Standard deviation	4,171	4,339	4,435	4,984	3,458	3,533
Log-normality test (P value)†	0.24	0.19	0.16	0.23	0.39	0.45
Panel B. Regression of k_0w_h on household attributes						
Number of sons	−781.3	−894.3	−859.2	−1013.1	−631.7	−693.0
	[303.6]**	[316.6]***	[326.8]***	[371.5]***	[276.7]**	[279.3]**
Number of daughters	−671.9	−805.2	−762.3	−787.8	−441.2	−471.2
	[354.7]*	[348.1]**	[372.2]**	[426.2]*	[259.0]*	[269.7]*
Father's years of schooling	24.6	2.3	19.6	17.5	19.1	23.6
	[89.6]	[91.0]	[93.2]	[109.4]	[74.8]	[77.1]
Value of houses	0.1	0.1	0.1	0.2	0.1	0.1
	[0.0]***	[0.1]**	[0.1]***	[0.1]***	[0.0]**	[0.0]**
Value of Agricultural equipment	0.3	0.3	0.3	0.3	0.3	0.3
	[0.1]***	[0.1]***	[0.1]***	[0.1]***	[0.1]***	[0.1]***
Fengrun	282.6	396.4	351.1	501.8	344.0	258.9
	[845.4]	[883.1]	[923.5]	[992.7]	[692.6]	[698.3]
Chicheng	−946.1	−979.1	−1042.2	−1127.9	−436.4	−530.9
	[806.1]	[843.7]	[878.6]	[967.0]	[656.0]	[671.8]
R^2	0.40	0.39	0.37	0.38	0.35	0.36

This table examines properties of the consumption level for the hypothetical son in each household, i.e., k_0w_h in Eq. (15) when transfers include only the bride-price. Results are estimated using the full multiple-son sample (293 sons in 140 households). The column number above each column in parenthesis refers to the corresponding specification in panel A of Table 4 used to generate the estimate of k_0w_h. Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

† P value of the log-normality test. For large P value, we cannot reject the hypothesis "distribution of the variable follows log normal"

Results are reported in panel B of Table 4. Consistent with our conjecture, taking into account the value of inherited land substantially increases the absolute value of the coefficients on schooling expenditures in all specifications. Our preferred specification in column (8) now suggests β=0.47 rather than 0.22.[31] That is, the estimated compensation coefficient doubles when we use a larger fraction of the lifetime parental transfer as a dependent variable. Using the same bootstrap procedure as described in Section 4.1,

[31] We also replicate panel A of Table 4 using households with land divisions. The compensation coefficient in the regression with full controls is 0.22 as well.

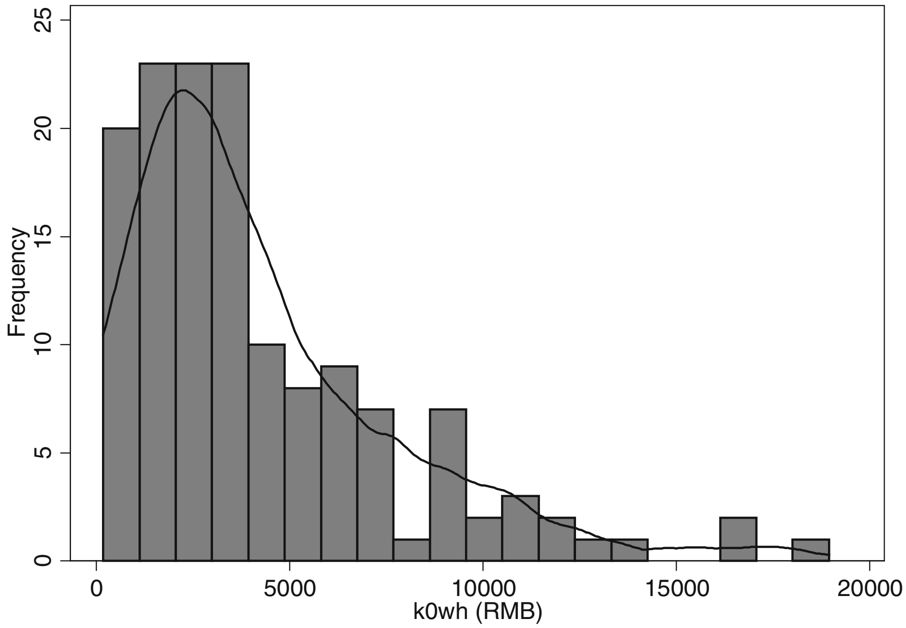

Fig. 4 Distribution of $k_0 w_h$ estimated from specification (8) in Panel A, Table 4. The curved solid line is the Kernel density

results in panel C indicate that the increase in β as we increase α is statistically significant at the 5 % level.

Qualitatively, both our multiplicative fixed-effects model and the log-linear model yield compensation coefficients that are increasing in α. However, our behavioral log-linear model does not allow for such dependence. Moreover, our log-linear derivation assumes μ is close to 1. Using $\beta=\alpha(1-\mu)r$, we can solve for μ as:

$$\mu = 1- \frac{\beta}{\alpha r}$$

Since marital transfers and land divisions represent a significant portion of lifetime transfers, let a conservative lower-bound estimate of α be 0.75. r is 1 plus the rate of return to schooling expenditure compounded up to the date of the wedding. Let a lower bound of r be 2.[32] Then, the 0.47 estimate of β implies that a lower bound estimate of μ is 0.69. An upper bound estimate of α is 0.9. Let an upper bound estimate of r be 3. Then an upper bound estimate of μ is 0.83. So we bound the estimate of μ, between 0.69 and 0.83.

With $0.69\leq\mu\leq0.83$, we can reject $\mu=0$. That is, holding a household's wealth constant, a son's lifetime consumption increases as his lifetime labor earnings increases. We also reject Becker's benchmark unitary model with equal concerns. As discussed earlier, we cannot reject the unitary model when parents do not have equal concerns for their children.

[32] Children in our sample have on average 8 years of schooling, and get married at the age of 23. Assuming that they start school at age 7, this implies that 8 years pass after they complete school before they get married. A back-of-the envelope calculation of the return to schooling at the time of marriage will be $r=(1+i)^8$, where i is the annual marginal return on education expenditure. Assuming that $0.09<r<0.15$, then $2<r<3$.

We can also reject a value of μ close to unity. Our estimate of μ provides evidence against a necessary assumption for our behavioral log linear model to be valid.

4.4 Heterogenous schooling coefficients

So far, we have maintained the assumption that both r and μ are homogenous across siblings. We now relax this assumption and discuss how that will affect the interpretation of our results. To fix ideas, assume that the heterogeneity in β takes the form: $(1-\mu_{ih})r_{ih}=(1-\mu)r+\upsilon_{ih}$ with $\mathrm{cov}(\upsilon_{ih}, a_{ih})=0$.

Consider the following regression:

$$\alpha t_{ih} = (1 + D_{ih}\delta)\omega_h - \beta s_{ih} + \gamma D_{ih} s_{ih} + \varepsilon_{ih} \tag{18}$$

where we interact schooling investments with children observables (D_{ih}). The coefficient γ captures heterogeneity in $\alpha(1-\mu)r$, which comes from sibling differences in either marginal returns in schooling investments (r), or the proportion of the schooling returns they retain for themselves (μ).

Table 6 reports the results using the bride-price as the dependent variable. For each specification in columns 1 to 6, we interact the same attributes with schooling investments as used to control for the parental bias.[33] Table 7 reports the same set of results, but using the bride-price plus land value as the dependent variable. The first thing to notice is that our estimates of the γs in Eq. (18) are very imprecise. The large standard errors are likely due to our relatively small sample size. Second, because of the weak explanatory power of the interaction terms,[34] adding them to the regressions does not significantly alter our estimates of β and δs relative to the specification without interaction terms (see Table 4).

Among all of the attributes interacted with educational expenditure, only the first son dummy, which is consistently positive across Tables 6 and 7, has a significant interaction effect. This means that the compensation coefficient $(1-\mu)r$ is smaller for the first sons, either due to their smaller r, or due to their larger μ.[35] The former is more likely to be the case for two reasons. First, first sons are more likely to work on the farm, thereby lowering returns to schooling investment (r) relative to their siblings. When schooling investments are efficient, lower returns to schooling leads to lower (optimal) investments in schooling. Our data show that this is the case: the first sons tend to receive lower schooling investments (1,386 yuan on average) than their siblings (1,601 yuan on average). Second, the focus on agricultural activities by first sons makes their return to education more likely to be shared by their parents. Therefore, if there is any difference, the first sons should have lower, not higher μ than their brothers.

In summary, we find evidence suggesting that $(1-\mu_{ih})r_{ih}\neq(1-\mu)r$. However the large standard errors preclude us from being able to make precise quantitative statements on the heterogeneity in marginal compensation.

[33] We also try specifications where all of the attributes are interacted with the schooling investments regardless of the way we control for k_{ih}. The conclusions are essentially the same.

[34] Comparing Table 6 with panel A of Table 4, and Table 7 with panel B of Table 4, we can see that they have similar R^2 for all specifications.

[35] Appendix D provides an illustrative explanation on why first sons can have both lower k_{ih} and lower $(1-\mu_{ih})r_{ih}$ at the same time.

Table 6 Multiplicative specification with interaction effects (bride-price as the dependent variable)

	(1)	(2)	(3)	(4)	(5)	(6)
Edu_exp ($-\beta_{t_1}$)	−0.32 [0.13]**	−0.32 [0.12]***	−0.28 [0.14]*	−0.24 [0.09]***	−0.45 [0.12]***	−0.32 [0.17]*
δ						
D_Height	0.26 [0.05]***					0.32 [0.08]***
D_Ag_Experience		0.02 [0.1]				−0.01 [0.13]
D_Nonag_Experience			0.10 [0.08]			0.05 [0.09]
D_Live_with_Parents				0.43 [0.12]***		0.49 [0.13]***
D_1st_Son					−0.27 [0.07]***	−0.29 [0.09]***
Interaction with edu_exp (γ)						
D_Height	−0.08 [0.2]					−0.08 [0.17]
D_Ag_Experience		1.10 [0.81]				0.18 [0.78]
D_Nonag_Experience			−0.32 [0.19]*			−0.04 [0.21]
D_Live_with_Parents				0.64 [0.64]		0.44 [0.48]
D_1st_Son					0.27 [0.18]	0.27 [0.19]
R^2	0.77	0.77	0.76	0.79	0.78	0.82
F test 1[†]						0.00
F test 2[‡]						0.51

Results are estimated using the full multiple-son sample (293 sons in 140 households). Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

[†] P value of the F test for the joint significance of k_{ih}

[‡] P value of the F test for the joint significance of the interaction terms

5 Robustness checks

5.1 Sensitivity to discount rates

The framework discussed in Section 2 is a static model while actual intra-household transfers are made over time and in different periods for different children. To make these transfers comparable in present value terms, all transfers made in each household must be discounted at rate d to a fixed point in time, which we set to be the year in which the first son was married. The discount rates represent the opportunity cost in

Table 7 Multiplicative specification with interaction effects (bride-price plus value of land as the dependent variable)

	(1)	(2)	(3)	(4)	(5)	(6)
Edu_exp	−0.45	−0.37	−0.32	−0.35	−0.68	−0.63
($-\beta_{t_3}$)	[0.17]***	[0.17]**	[0.18]*	[0.15]**	[0.22]***	[0.27]**
δ						
D_Height	0.02					0.10
	[0.07]					[0.12]
D_Ag_Experience		−0.07				−0.05
		[0.11]				[0.16]
D_Nonag_Experience			0.10			0.00
			[0.07]			[0.13]
D_Live_with_Parents				0.09		0.22
				[0.12]		[0.16]
D_1st_Son					−0.33	−0.41
					[0.06]***	[0.12]***
Interaction with edu_exp (γ)						
D_Height	0.21					0.16
	[0.22]					[0.32]
D_Ag_Experience		1.38				0.57
		[1.15]				[1.19]
D_Nonag_Experience			−0.33			0.05
			[0.21]			[0.3]
D_Live_with_Parents				1.01		0.86
				[0.89]		[0.7]
D_1st_Son					0.45	0.52
					[0.26]*	[0.3]*
R^2	0.77	0.77	0.76	0.78	0.82	0.85
F test 1†						0.00
F test 2‡						0.29

Results are estimated using the multiple-son sample with land divisions (186 sons in 89 households). Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

† P value of the F test for the joint significance of k_{ih}

‡ P value of the F test for the joint significance of the interaction terms

children's consumption when transfers are provided 1 year earlier; its value should reflect the real return to saving in rural China. In light of the low nominal deposit rates on savings in China's formal financial institutions, we select relatively low values for d of 0.025 and 0.05.[36] In columns (1)–(6) of panel A, Table 8, we compare different

[36] Deposit rates on savings are reported in the annual Zhongguo Jinrong Nianjian (Almanac of China's Finance and Banking).

Table 8 Multiplicative specification with discounting

	Basic			Preferred specification		
	d=0	d=0.025	d=0.05	d=0	d=0.025	d=0.05
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Bride-price as dependent variable						
Edu_exp	−0.34	−0.28	−0.20	−0.22	−0.18	−0.15
$(-\beta_{t_1})$	[0.17]*	[0.1]***	[0.09]**	[0.1]**	[0.08]**	[0.07]**
δ						
D_Height				0.31	0.31	0.31
				[0.09]***	[0.09]***	[0.09]***
D_Ag_Experience				0.00	−0.02	−0.04
				[0.11]	[0.1]	[0.1]
D_Nonag_Experience				0.05	0.05	0.05
				[0.1]	[0.1]	[0.1]
D_Live_with_Parents				0.60	0.61	0.61
				[0.16]***	[0.16]***	[0.15]***
D_1st_Son				−0.23	−0.05	0.14
				[0.09]**	[0.09]	[0.09]
R^2	0.75	0.77	0.78	0.82	0.81	0.81
F test 1[†]				0.00	0.00	0.00
Panel B. Bride-price + land division as dependent variable						
	(1)	(2)	(3)	(4)	(5)	(6)
Edu_exp	−0.67	−0.57	−0.47	−0.47	−0.42	−0.38
$(-\beta_{t_3})$	[0.14]***	[0.19]***	[0.19]**	[0.2]**	[0.2]**	[0.19]*
δ						
D_Height				0.19	0.21	0.23
				[0.35]	[0.34]	[0.34]
D_Ag_Experience				1.40	1.12	0.92
				[1.99]	[1.55]	[1.18]
D_Nonag_Experience				−0.25	−0.17	−0.14
				[0.42]	[0.4]	[0.4]
D_Live_with_Parents				1.00	0.70	0.44
				[0.98]	[0.84]	[0.76]
D_1st_Son				0.29	0.26	0.21
				[0.37]	[0.37]	[0.37]
R^2	0.81	0.83	0.84	0.84	0.84	0.85
F test 1[†]				0.01	0.23	0.46
Panel C. Bootstrap results						
Mean $(\beta_{t_3}/\beta_{t_1})$	1.86	1.72	1.10	2.99	2.02	1.58
Prob $(\beta_{t_3}/\beta_{t_1} > 1)$	0.99	1.00	0.99	0.97	0.97	0.96

Panel A reports results using the full multiple-son sample (293 sons in 140 households)

Panel B reports results using the multiple-son sample with land divisions (186 sons in 89 households)

Panel C reports the mean of β_{t_3}/β_{t_1} and the probability of $\beta_{t_3}/\beta_{t_1} > 1$ among 100 bootstrapped samples. Each sample draws 89 households with replacement using the original son sample with household divisions

Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

[†] P value of the F test for the joint significance of k_{ih}

specifications—the basic ones assuming "equal concerns" and ones with the full set of controls—under different discount rates in order to test for the sensitivity of our results.

Two patterns are notable in panel A. First, estimated coefficients on schooling investments decline in magnitude as the discount rate increases. Second, estimates of δ do not vary across alternative discount rates either in terms of their magnitudes or standard errors, except for a time related factor: the first son indicator. Coefficients on the first son dummy (row 6) consistently increase with discount rates, and become positive but insignificant under reasonable discounting rates (d=0.025, 0.05). We therefore conclude that first sons do not receive significantly higher consumption than other siblings.

These patterns are due to the correction for discounting. Educational expenditures occur before marital transfers. If we discount everything to the time when the first son got married,[37] the differences in education expenditure across siblings by age will widen. Meanwhile, the difference in marital transfers by age will be compressed. So discounting increases the value of schooling expenditure of the first son and reduces the value of marital transfers to the second son. This logic explains why the estimates of the first son affects increases with the discount rate.

Panel B includes the same set of results but using bride-price plus land division as the dependent variable ($t_3=t_1+t_2$). Even with proper discounting, the magnitude of β_{t_3} is typically double that of β_{t_1} for most of the specifications. On the basis of the bootstrap results in panel C, we cannot reject the null hypothesis in Eq. (14).

5.2 Measurement error in schooling investments

Given that our education expenditure variable mainly comes from recall data, measurement error might exist, which biases estimates of β towards zero (attenuation bias). One way to correct this bias is to adopt an instrumental variable approach. "Years of schooling" is an ideal candidate to serve as an instrument for education expenditure: First, it is positively correlated with education expenditure; and second, its measurement error, if there is any, is not likely to be correlated with that of education expenditure.

In panel A of Tables 9 and 10, we replicate all specifications in Table 8 with generalized method of moments (GMM) estimation, using bride-price and bride-price plus land value as the dependent variable, respectively. The "first stage" regressions and reduced form regressions are presented in Panel B and Panel C, respectively. As expected, "years of schooling" is a good predictor of education expenditure in the first stage. (1 year of schooling costs about 700 yuan on average.) Comparing the OLS with the IV results (panel A in Table 9 versus panel A in Table 8, and panel A in Table 10 versus panel B in Table 8), both the estimates of the β and δ do not change much, although the standard error of the former increases, thereby reducing its significance. Thus, there is no evidence of a strong quantitative attenuation bias due to measurement error in schooling expenditure.[38]

[37] Choosing a different baseline for the discounting will generate the same trend.

[38] We cannot reject the hypothesis of some attenuation bias if measurement errors in schooling expenditure and years of schooling are positively correlated.

Table 9 IV version of multiplicative model, Table 8, panel A (bride-price as the dependent variable)

Panel A. GMM estimates						
	d=0	d=0.025	d=0.05	d=0	d=0.025	d=0.05
Edu_exp	−0.34	−0.30	−0.25	−0.16	−0.16	−0.16
($-\beta_{t_1}$)	[0.18]*	[0.15]**	[0.12]**	[0.16]	[0.14]	[0.12]
δ						
D_Height	0.21	0.19	0.18	0.27	0.29	0.31
	[0.08]***	[0.08]**	[0.09]**	[0.09]***	[0.09]***	[0.09]***
D_Ag_Experience				−0.01	−0.04	−0.07
				[0.1]	[0.09]	[0.09]
D_Nonag_Experience				0.02	0.01	0.00
				[0.09]	[0.09]	[0.09]
D_Live_with_Parents				0.49	0.50	0.52
				[0.14]***	[0.14]***	[0.14]***
D_1st_Son				−0.20	−0.06	0.11
				[0.09]	[0.09]	[0.09]**
R^2	0.78	0.79	0.80	0.82	0.82	0.82
F test 1				0.00	0.00	0.00
Panel B. 1st stage						
yr_sch	689.56	738.59	797.36	658.51	697.84	745.68
	[158.2]***	[177.41]***	[202.25]***	[160.05]***	[179.62]***	[205.33]***
R^2	0.72	0.72	0.71	0.74	0.73	0.72
Panel C. Reduced form						
yr_sch	−95.51	−140.22	−155.77	−32.28	−73.10	−100.80
	[214.73]	[183.12]	[158.77]	[138.28]	[153.97]	[159.03]
R^2	0.78	0.79	0.81	0.82	0.82	0.82

This table replicates panel A of Table 8, while using years of schooling as an instrument for educational expenditure. Results are estimated using full the multiple-son sample (293 sons in 140 households)

Panel A reports the GMM estimates using the value of the bride-price as the dependent variable. Panel B reports the coefficient on years of schooling in first stage where the dependent variable is educational expenditure. Panel C reports the multiplicative fixed effects specification where educational expenditure is replaced by years of schooling

Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

† P value of the F test for the joint significance of k_{ih}

6 Conclusion

This paper studied how traditional agricultural families in Hebei, China, compensated their sons for unequal schooling expenditures. Our empirical strategy is different from most other papers on inter vivos transfers from parents to children. In the literature, researchers usually only observe transfers within a fixed short calendar window (e.g., transfers in the previous year), a setting in which many children receive zero transfers. Instead, we focus on transfers tied to important family lifecycle events, which provides a more comprehensive measure of family transfers.

Table 10 IV version of multiplicative model, Table 8, panel B (bride-price plus value of land as the dependent variable)

Panel A. GMM estimates						
	d=0	d=0.025	d=0.05	d=0	d=0.025	d=0.05
edu_exp ($-\beta_{t_3}$)	−0.75	−0.62	−0.52	−0.47	−0.43	−0.39
	[0.23]***	[0.19]***	[0.17]***	[0.26]*	[0.23]*	[0.21]*
δ						
D_Height	0.21	0.19	0.18	0.03	0.03	0.03
	[0.08]***	[0.08]**	[0.09]**	[0.07]	[0.07]	[0.07]
D_Ag_Experience				0.05	0.04	0.03
				[0.11]	[0.11]	[0.1]
D_Nonag_Experience				0.05	0.06	0.08
				[0.1]	[0.1]	[0.1]
D_Live_with_Parents				0.21	0.22	0.24
				[0.12]*	[0.13]*	[0.13]*
D_1st_Son				−0.19	−0.09	0.03
				[0.08]**	[0.08]	[0.08]
R^2	0.81	0.83	0.84	0.84	0.84	0.85
F test 1				0.00	0.03	0.04
Panel B: 1st stage						
yr_sch	676.46	713.15	760.30	596.09	618.56	649.84
	[155.57]***	[165.44]***	[181.14]***	[136.53]***	[138.01]***	[145.13]***
R^2	0.72	0.72	0.72	0.75	0.76	0.76

Table 10 (continued)

Panel C. Reduced form						
Yr_sch	−506.06	−448.86	−382.53	−276.01	−307.77	−309.07
	[245.35]**	[212.94]**	[193.23]**	[220.44]	[206.47]	[189.54]
R^2	0.79	0.81	0.83	0.83	0.83	0.83

This table replicates panel B of Table 8, while using years of schooling as an instrument for educational expenditure. Results are estimated using the multiple-son sample with land divisions (186 sons in 89 households)

Panel A reports the GMM estimates using the value of bride-price plus land division as the dependent variable is. Panel B reports the coefficient on years of schooling in the first stage where the dependent variable is educational expenditure. Panel C reports the multiplicative fixed effects specification where educational expenditure is replaced by years of schooling

Robust standard errors are in brackets

*Significant at 10 %

**Significant at 5 %

***Significant at 1 %

† P value of the F test for the joint significance of k_{ih}

We estimate two empirical models of parental transfers, the log-linear family fixed effects model and the multiplicative family fixed effects model. For both theoretical and empirical reasons, we reject the log-linear model as a behavioral model of how these rural families behave. Based on the estimates of the multiplicative family fixed effects model, we conclude that the marginal compensation coefficient is positive and quantitatively large but smaller than 1. Since marital transfers and land division together are a substantial fraction of total parental transfers, it is likely that that sons with more schooling have higher lifetime consumption.

Our preferred estimate of the marginal compensation coefficient is 0.47. The marginal compensation coefficient is a product of the fraction of lifetime transfers observed, α, the fraction of labor earnings of the child which is appropriated by the family, $1-\mu$, and the gross return to schooling investment, r. Since marital transfers and land divisions together imply that α is greater than 0.8, and the gross return to schooling has to be larger than 1, the fraction of labor earnings of the son which he retains is likely significantly less than 1. This fraction accords well with descriptive evidence in which rural households pool their earnings.

There are a few directions for further research. First, it will be important to study the salience of the log-linear model of parental transfers in settings where the children's incomes are not pooled with their parents and μ is closer to unity, a necessary condition for the log-linear behavior model to be valid. Second, as μ is greater than zero, our model suggest that parental investments in children may deviate from the efficient level due to strategic considerations (as shown by Park 2003, for example), which deserves future empirical investigations. Third, our results suggest that it may be also useful in other environments to retrospectively study major transfers by parents to their children. Fourth, it is important to study the determinants of inter vivos transfers to daughters. This requires dealing with selection issues that arise because a substantial number of daughters do not receive marital transfer from their families. And fifth, as discussed earlier, the sample size for our study is not large. Any inference which we draw must take that into consideration.

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Appendix

A. A framework for parental investments and transfers

This appendix provides a general framework of intra-household resource allocation. Decisions are made in two stages. In the first stage, parents choose schooling investments (s_{ih}) among different children; in the second stage, consumption is allocated among family members.

We start from the second stage of the problem. Taking s_{ih} as given, consumption levels are determined by solving (P1):

$$\max_{c_h, c_{1h}, c_{2h}} U_h(c_h, c_{1h}, c_{2h}) = (1 - k_{1h} - k_{2h})u(c_h) + k_{1h}u(c_{1h}) + k_{2h}u(c_{2h}) \quad \text{(P1)}$$

subject to the household budget constraint:

$$c_h + c_{1h} + c_{2h} = w_h(s_{1h}, s_{2h}) = m_h + \sum_{i=1}^{2}(R(s_{ih}) - C(s_{ih}, a_{ih}, m_h)) \quad (19)$$

It is convenient to impose a homothetic assumption on $U_h(\cdot)$, say $u(c)=\ln(c)$ (see, Becker (1991)). Then optimal consumption levels are proportional to the total family wealth:

$$c^*_{1h} = k_{1h}w_h \quad (20)$$

$$c^*_{2h} = k_{2h}w_h \quad (21)$$

$$c^*_h = (1 - k_{1h} - k_{2h})w_h \quad (22)$$

This consumption allocation solution can be rationalized through either a unitary or collective model. Their difference lies in the interpretation of k_{ih}. In the context of the unitary model, equation (P1) represents the utility function of altruistic parents, who allocate consumption among family members. k_{ih} can be interpreted as the relative weight they put on their children's welfare relative to their own. Under Chiappori's (1988, 1992) collective model, family members bargain efficiently over the division of family wealth to obtain their own consumption. The efficient bargaining results can be implemented through maximizing a "household welfare" function as in (P1), where k_{ih} represents an individual's bargaining power. In general,

$$k_{ih} = k_{0h} + \mu\frac{rs_{ih}}{w_h}; \mu \geq 0 \quad (23)$$

When $\mu>0$, the parents' marginal utility from a child's consumption is increasing in the child's earnings from schooling expenditure relative to family wealth.

In the first stage, parents make decisions on schooling to maximize their own utility function (assumed to be homothetic as well), while taking into account the allocation decisions (20)–(22):

$$\begin{aligned} &\max_{c_h, c_{1h}, c_{2h}} (1 - b_{1h} - b_{2h})\ln c_h + b_{1h}\ln c_{1h} + b_{2h}\ln c_{2h} \\ &\text{s.t. } (20) - (22) \end{aligned} \quad \text{(P2)}$$

where b_{ih} are the relative weights they put on their children's welfare relative to their own.

Notice that our framework is equivalent to Becker's benchmark unitary model when $b_{ih}= k_{ih}$ and $\mu=0$.

The value of μ is crucial for the efficiency of schooling investments. When $\mu=0$, k_{ih} is independent of schooling attainment. Optimal schooling investments $s^*_{ih}(a_{1h}, a_{2h}, m_h)$ satisfy:

$$\frac{\partial w(s_{1h}, s_{2h})}{\partial s_{ih}} = 0$$
$$\text{i.e.} \frac{\partial R(s_{ih})}{\partial s_{ih}} = \frac{\partial C(s_{ih}, a_{ih}, m_h)}{\partial s_{ih}}$$

i.e., the *efficient schooling investment level* (s_i^*) will maximize total family wealth.[39] This is true under both the unitary and collective model. Even when household members bargain over the division of family wealth and parents have their own private interests, schooling investment is chosen to maximize total family wealth. Although parental preferences may differ from the bargaining weights, ($b_{ih} \neq k_{ih}$), the best the parents can do given the allocation constraints (20)–(22) is to maximize total family wealth. This result is reminiscent of Becker's Rotten Kid Theorem.

When $\mu > 0$, schooling investment will not be wealth maximizing. Under the unitary model, $b_i(s_i) = k_i(s_i)$. The optimal solution for child *i's* schooling s'_{ih} satisfies the first-order condition of (P2):

$$\frac{\partial w(s_{1h}, s_{2h})}{\partial s_{ih}} = -\mu r \cdot \ln \frac{k_{ih}}{1 - k_{1h} - k_{2h}}$$

When $k_{ih} > 1 - k_{1h} - k_{2h}, \frac{\partial w_h}{\partial s_{ih}} < 0$, hence $s'_{ih} > s^*_{ih}$; when $k_{ih} < 1 - k_{1h} - k_{2h}, \frac{\partial w_h}{\partial s_{ih}} > 0$, hence $s'_{ih} < s^*_{ih}$, i.e., parents will *over/under* invest (relative to the wealth maximizing level) in one child's schooling if he receives a higher/lower weight in the parental utility function than do the parents.[40]

Under the collective model, $b_{ih}(s_{ih}) \neq k_{ih}(s_{ih})$. The optimal solution for child *i's* schooling s''_{ih} satisfies:

$$\frac{\partial w(s_{1h}, s_{2h})}{\partial s_{ih}} = -\mu r \cdot \frac{b_{ih}(1 - k_{jh}) - k_{ih}(1 - b_{jh})}{k_{ih}(1 - k_{1h} - k_{2h})}$$

In particular, if $b_{ih} < k_{ih}(s''_{ih})$ for $i = 1, 2$, then $\frac{\partial w_h}{\partial s_{ih}} > 0$ and therefore $s''_{ih} < s^*_{ih}$. This means that, if at the wealth maximizing schooling level, children's bargaining powers are greater than their parents would like them to be, the parents will invest less in children's schooling relative to the efficient level.

B. Estimation of the multiplicative model

Given a random sample of households $h = 1, \ldots, H$, each with 2 children $i = 1, 2$, consider the regression:

$$y_{ih} = (1 + D_{ih}\delta) \cdot \omega_h + X_{ih}b + \varepsilon_{ih} \quad (24)$$

where y_{ih} is the dependent variable; D_{ih} and X_{ih} are vectors of children's characteristics; and ω_h is the household fixed effect.[41] The parameters to be identified are $\theta = (\delta, \{\omega_h\}_{h=1}^{H}, b)$.

A nonlinear least squares (NLS) estimator of $\theta, \widehat{\theta}$, solves:

$$\min_{\theta} \sum_{h=1}^{H} \sum_{i=1}^{2} [y_{ih} - (1 + D_{ih}\delta) \cdot W_h - X_{ih}b]^2$$

[39] Recent empirical evidence like Fitzsimons and Malde (2014) shows that e¢ ceincy in childrenís education investment does not depend on number of children in the family.

[40] Efficiency is reached when all family members recieve equal weights: $k_1 = k_2 = 1 - k_1 - k_2 = 1/3$.

[41] Equation (24) is a generalized version of (17) and (18).

To reduce the computational burden of a global search for all parameters, we take advantage of the partial linearity in (24): given a value of δ, δ_0, equation (24) is a linear model. $\{\omega_h\}_{h=1}^{H}$ and b can be estimated through OLS. Denoting these estimators as $\{\omega_h^{ols}(\delta_0|X_{ih},D_{ih})\}_{h=1}^{H}$ and $b^{ols}(\delta_0|X_{ih},D_{ih})$, our problem is to find:

$$\widehat{\delta} = \arg\min_{\delta_0} \sum_{h=1}^{H} \sum_{i=1}^{2} \left(y_{ih} - (1 + D_{ih}\delta_0)\cdot\omega_h^{ols}(\delta_0|X_{ih}, D_{ih}) - X_{ih}b^{ols}(\delta_0|X_{ih}, D_{ih})\right)^2$$

Given $\widehat{\delta}$, the final estimator $\widehat{\theta}$ is:

$$\widehat{\theta} = \left(\widehat{\delta}, \left\{\omega_h^{ols}\left(\widehat{\delta}|X_{ih}, D_{ih}\right)\right\}_{h=1}^{H}, b^{ols}\left(\widehat{\delta}|X_{ih}, D_{ih}\right)\right)$$

C. The incidental parameter problem

The least square estimators outlined above are subjected to a so-called "incidental parameter" problem. As first observed by Neyman and Scott (1948), standard estimators of nonlinear panel data models are usually inconsistent if the length of the panel is small relative to the number of observations. In this case, the finite sample bias in the fixed effects parameters ($\{\omega_h\}_{h=1}^{H}$ in our context) will contaminate estimates of other parameters (δ and b in (24)).

Given the partial nonlinearity feature of our model, an alternative approach that can help get around the incidental parameter problem is quasi-differencing. To see this, divide (24) by $1+D_{ih}\delta$ on both sides, and then take the sibling difference within the same household to obtain:

$$\Delta \frac{y}{1 + D\delta}_h = \Delta \frac{X}{1 + D\delta}_h b + \upsilon_h \tag{25}$$

where $\Delta \frac{x}{1+D\delta}_h = \frac{x_{1h}}{1+D_{1h}\delta} - \frac{x_{2h}}{1+D_{2h}\delta}$ and $\upsilon_h = \frac{\varepsilon_{1h}}{1+D_{1h}\delta} - \frac{\varepsilon_{2h}}{1+D_{2h}\delta}$. Notice that household fixed effects are eliminated in (25). The parameters to be determined are (δ, b), which can be consistently estimated through NLS.[42]

In order to compare the accuracy and efficiency between our least-square estimate and the above quasi-differencing estimate, we conduct Monte Carlo experiments. The number of households is 140, consistent with our multiple-sons sample. In each simulation, each household is endowed with a k_0w_h, which is assumed to be log normal distributed. Its distribution parameters are calibrated from the moments of the household fixed effects $\{w_h\}_{h=1}^{H}$ that we estimate from the data.

Each household has two children, one of whom is the "first son". We randomly generate two other attributes of the children, "indicator of the taller son" (D_{ih}^{taller}) and "indicator of living with parents post marriage" (D_{ih}^{livep}), such that their variance-covariance matrices with the first son indicator (D_{ih}^{1stson}) replicate the ones in the actual data.

[42] For consistency we need $\mathrm{cov}\left(\Delta \frac{X}{1+D\delta}_h, \upsilon_h\right) = 0$ in (25). This condition is ensured under the efficient schooling investment assumption.

We regress schooling expenditure on $k_0 w_h$ and children's attributes using the actual dataset. We then use the coefficients and the distribution parameters of the error terms to generate the simulated schooling investments. Note that the error terms here serve as the ability endowments of the children. The intervivos transfers are simulated using equation (17) with error terms randomly generated to match the sample moments of the marital transfers (t_{ih}). The resulting data set is $\{s_{ih}, t_{ih}, D_{ih}^{taller}, D_{ih}^{livep}, D_{ih}^{1stson}\}$. We estimate the model:

$$t_{ih} = \left(1 + \delta_1 D_{ih}^{taller} + \delta_2 D_{ih}^{livep} + \delta_3 D_{ih}^{1stson}\right) \cdot \omega_h + \beta s_{ih} + \varepsilon_{ih}$$

with the "true" parameters being set as $(\beta, \delta_1, \delta_2, \delta_3) = (-0.3, 0.2, 0.4, -0.2)$. Note that we allow $\mathrm{cov}(s_{ih}, \omega_h) > 0$ and $\mathrm{cov}(s_{ih}, D_{ih}) > 0$, but keep $\mathrm{cov}(s_{ih}, \varepsilon_{ih}) = 0$ in the simulated data, which are the basic identification assumptions maintained throughout our paper.

We conduct a Monte Carlo analysis with 500 simulations. The results are reported in an appendix table available online. Our results show that, the least square approach used in our paper delivers highly precise estimates. By contrast, the quasi-differencing results exhibit larger bias and larger standard errors, especially when all three attributes of the children are controlled for simultaneously. [43]We conclude that, even though our least square estimates may be theoretically inconsistent, their finite sample bias is negligible and they are more efficient than the consistent quasi-differencing estimates. We therefore base our inferences on the least squared results in our paper.

D. An illustrative explanation for why the first son can have both smaller k_{ih} and smaller $(1-\mu_{ih})r_{ih}$ at the same time

In an appendix figure available online, we show, in an illustrative way, the fitted linear relationship between educational expenditure and the transfers for the first son and his sibling (denoted as the second son).[44] The readers are referred to that figure for the discussion below. By definition, the two lines in the figure should pass through the sample mean for the two sons, denoted by point A and B, respectively. Their relative positions in the figure are due to the fact that, on average, eldest sons receive less schooling investments and intervivos transfers than their siblings in the data.[45]

Subfigure (1) illustrates the case of Table 4, where the slopes of the two fitted lines, i.e. the marginal compensation coefficients, are assumed to be the same. But intercepts of the lines can be different, which identifies δ in equation (17) for the first son. The assumption that the two lines are parallel, along with the relative position of point A and B, determines that l_1 lies below l_2, which is consistent with the negative value for δ for the first son in Table 4.

[43] This conclusion is similar to Greene (2004). Using Monte Carlo methods, he finds that the coefficients in the Tobit model with fixed effects are "unaffected" by the incidental parameter problem. He observes that the estimators' behavior in models with a continuous dependent variable (whether truncated or not) are quite different from binary choice models.

[44] For simplicity, we assume that $w_h = w$ for all h.

[45] On average, educational expenditure, brideprice, and brideprice plus land for the first son are 1386 yuan, 4440 yuan, and 6492 yuan, respectively. Mean value for the same set of variables for the other sons are 1601 yuan, 5299 yuan, and 7697 yuan, respectively.

Subfigure (2) illustrates the case of Table 6, where the slopes of the two lines are allowed to be different. Notice that the coefficient on the interaction between the first son dummy with the educational expenditure in Table 6 is positive, which means that the marginal coefficient is smaller for the first son. This result, along with the relative position of A and B, determines that magnitude of δ for the first son should be larger in Table 6 than in Table 4, which is consistent with our results.

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