Magical Transition? Intergenerational Educational and Occupational Mobility in Rural China: 1988-2002 [1]

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ABSTRACT

This paper presents evidence on intergenerational educational and occupational mobility in Rural China over a period of 14 years (1988-2002). To understand whether the estimated intergenerational persistence can be driven solely by unobserved heterogeneity, we implement biprobit sensitivity analysis (Altonji et al. (2005)) and heteroskedasticity based identification of Klein and Vella (2009). The empirical results show that there have been dramatic improvements in occupational mobility from agriculture to non-farm occupations; a farmer's children are not any more likely to become farmers in 2002, even though there was significant persistence in occupation choices in 1988. In contrast, the intergenerational mobility in educational attainment has remained largely unchanged for daughters, and it has deteriorated significantly for sons. There is strong evidence of a causal effect of parental education on a son's schooling in 2002. We provide some possible explanations for the dramatic divergence between occupational and educational mobility in rural China from 1988 to 2002.

Key Words: Intergenerational Mobility, Rural China, Occupational Choice, Educational Attainment, Economic Reform, Heteroskedasticity Based Identification, Biprobit Sensitivity Analysis

JEL Classification: O12, J62

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(1) Introduction

Intergenerational transmission of economic status has been the focus of a growing literature in economics.[2] There is now a substantial literature on developed countries that shows significant persistence of economic status across generations; the estimated partial correlation between income of parents and children falls in the range of 0.3 to 0.6 (see Black and Devereux (2010), Blanden et al. (2005) and Solon (1999, 2002), Mazumder (2005)). In contrast, economic analysis of intergenerational mobility in the context of developing and transition countries remains a relatively little explored area of research; among the few available contributions are Lillard and Willis (1995) on Malaysia, Hertz (2001) on South Africa, Sato and Li (2007a, 2007b) on China, Emran and Shilpi (2011) on rural Nepal and Vietnam.[3]

In this paper, we analyze the evolution of intergenerational economic mobility in rural China from 1988 to 2002. The central question addressed in this paper is: if we compare two snapshots of rural China in 1988 and 2002, has Chinese rural society on an average become more or less economically mobile over a decade and a half during which deep and wide-ranging market oriented economic reforms have been implemented?[4] The empirical analysis focuses on educational and occupational persistence across generations.

China is an interesting and important case study for understanding the nature of economic mobility during transition from socialist command economy to a more market oriented one. While China achieved extraordinary growth and poverty reduction over the last few decades, the growing inequality has become a focal point of concern for policy makers. Although cross-sectional inequality and intergenerational mobility are different concepts, one would in general expect inequality in opportunities to be manifested as high cross-sectional inequality. But high cross-sectional inequality at a given point of time may not necessarily imply that inequality in opportunities is also high. For example, government policy interventions in education and health can improve mobility and such policies may be implemented partly in response to high cross-sectional inequal-

[2]See, for example, Arrow et al. (2000), Behrman and Rosenzweig (2002), Black and Devereux (2010), Black et al. (2005, 2007), Bjorklund et. al. (2006), Bjorklund and Salvanes (2010), Dearden et al. (1997), Mazumder (2005), Aaronson and Mazumder (2008), Hertz (2005), Hertz et al. (2007), Mulligan (1999), Solon (1999, 2002, 2004), Fields et al. (2005), Bowles et. al. (2005), Blanden et al. (2005), World Development Report (2005).

[3]Hertz et al. (2007) provide a analysis of the basic correlations in educational attainment in a sample of 42 countries.

[4]We thus do not focus on differences across cohorts of children as is common in the literature on trends in intergenerational mobility (see, for example, Aaronson and Mazumder (2008)).

ity. [5] Have the economic reforms in China that begun in 1978 such as household responsibility system, the gradual relaxation of Hukou system (restrictions on geographic mobility), and 9 years of compulsory schooling policy (starting from 1986) expanded equality of economic opportunities in rural China?[6] Or the market oriented reforms have instead generated inequality in economic opportunities, thus making the 'accident of birth' increasingly more important in shaping the opportunities faced by an individual?[7] If the same reforms that resulted in high growth with high cross-sectional inequality have also generated inequality in opportunity, it has profound economic and political implications. A related important issue is whether there are significant gender differences in intergenerational educational and occupational mobility in a given period and over time (over a period of 14 years). To the best of our knowledge, there is no rigorous analysis of the evolution of intergenerational educational and occupational mobility in rural China in the post reform period in existing economics literature.[8] We use two rounds of Chinese Household Income Project (CHIP) survey data for the years 1988 and 2002 for the analysis of intergenerational persistence in educational attainment and occupational choices over a span of almost a decade and a half. The data in both 1988 and 2002 are based on almost identical questionnaires (2002 round has some added information), and are comparable. This allows us to trace out the *changes* in intergenerational occupational and educational persistence from 1988 to 2002.

The economics literature on the intergenerational economic mobility in developed countries focuses on income mobility, especially between father and son.[9] There is also a relatively small literature on occupational mobility in labor economics, again mostly in the context of developed

[5] As argued by Friedman (1962), a given extent of cross-sectional inequality is a bigger concern in a rigid society where the pattern of inequality is self-reproducing compared to a mobile society where such inequality can decline in short to medium term because of equality of opportunity.

[6] Given the spectacular growth in income in rural China, one would expect a significant increase in parental investment in children's education, especially because the family may no longer be dependent on the income from child labor (luxury axiom of child labor a la Basu and Van (1998)). The growth in per capita income in rural China has been impressive after the initiation of economic reform in 1978. Per capita income in rural China grew from 133.6 yuan in 1978 to 544.9 yuan in 1988 to 2475.6 yuan in 2002 (National Bureau of Statistics of China).

[7] The existing Sociological literature on mobility in China find that economic mobility of the poor peasants were positively affected by the communist policies before the economic reform of 1978. See, for example, Cheng and Dai (1995).

[8] The only paper we are aware of is Sato and Li (2007a) which analyzes the educational mobility among three generations of men in rural China using 2002 CHIP data set. In contrast, we focus on the *changes*, if any, in the *average* economic mobility in rural China from 1988 to 2002. In a recent paper, Gong et al. (2010) analyze income mobility in *urban* China.

[9] The handful of studies focusing on the intergenerational correlations for daughters include Chadwick and Solon (2002) for USA and Dearden, Machin and Reed (1997) for UK.

countries.[10] Similar to income mobility, the existing literature on occupational mobility focuses primarily on the father-son linkage (see Lentz and Laband (1983), and Dunn and Holtz-Eakin (2000) on U.S, Sjogren (2000) on Sweden, and Behrman et. al. (2001) on Latin America). It is, however, difficult to focus on income or consumption mobility across generations in the context of most of the developing countries because of the unavailability of appropriate data. It is extremely difficult, if not impossible, to find reliable data on income and consumption of the parental generation. As emphasized in recent literature, to understand the intergenerational persistence in income, it is important to have data over many periods (years) so that the problem of measurement error due to transitory shocks can be addressed properly.[11] In the face of such data limitations, we use education and occupation as two salient indicators of economic status for which the requisite data are available.[12]

There are a number of different channels through which intergenerational linkages in occupation and education can and do operate (for recent discussions, see Bjorklund and Salvanes (2010), Solon (1999), Behrman et al. (2003), Bowels, Gintis, and Osborne Groves (2005), and for the pioneering analysis of intergenerational mobility, see Becker and Tomes (1976, 1979, and 1986)). Some of the channels are tangible, such as parental investment in children's education, bequests from parents and access to parent's social network.[13] However, a significant part of the intergenerational correlations presumably arises from the effects of intangible factors such as genetic transmissions of ability and preference from parents to children, learning externalities (learning at the 'dinner table' and learning by watching and informal apprenticeship in the case of occupation), role model effects and transfer of reputation capital (likely to be especially important in occupation choices). At least part of the intergenerational persistence in economic outcomes is likely to be due to the unobserved factors common across generations, and this poses challenge in understanding possible causal effects of parental education and occupation on their children's

[10] The intergenerational occupational mobility has, however, been the the focus of a large, mostly descriptive, literature in sociology.

[11] As shown by Mazumder (2005), the transitory shocks can be a lot more persistent than usually thought of, and if one relies on the average over a few years data, it might not be possible to tackle the measurement error adequately.

[12] Following Emran and Shilpi (2011), we focus on mobility out of agriculture to non-farm activities as a measure of occupational mobility in rural areas. The non-farm occupation is defined as non-agricultural rural occupation. There is substantial evidence that rural non-farm activities are avenues for poor households to escape from poverty traps. For a general discussion on rural nonfarm activities in developing countries, see Feder and Lanjou (2001).

[13] Bequests may relax credit constraint and thus induce children to start their own business in the non-farm sector.

choices. The distinction between the genetic and environmental influences on intergenerational linkages has been emphasized in the economic literature on intergenerational mobility, and it can be important from a policy perspective. If the observed intergenerational persistence is driven primarily by genetic transmissions of ability and preference across generations, the role for policy interventions in promoting economic mobility in a society may be more limited (Solon (1999, 2002), Bjorklund and Salvanes (2010)). In contrast, when environmental factors are important, there may be a wider scope for government policy interventions to alleviate persistent inequality in economic opportunities.[14]

Although the economic literature on intergenerational mobility has been fraught with the difficulties in addressing the selection on unobservable characteristics, especially genetic correlations, our study enjoys an important advantage in this regard. We are interested in understanding the possible *changes* in the intergenerational linkages over a relatively short period of time (1988-2002); thus the *direction of change* over time will be well identified under the plausible assumption that the genetic correlations do not change in any significant way in a span of 14 years. However, in the event that we find evidence of significant intergenerational linkages in a given year, we need to address the possibility that the estimated partial correlations between parents' and children's choices/outcomes are not causal, rather primarily driven by unobserved heterogeneity.

From the recent economic literature, one can identify three different approaches to uncovering causal influence of parental economic status on children's economic fortunes: (i) twin sample (see, for example, Behrman and Rosenzweig (2002)) (ii) adoptees sample (see, for example, Bjorklund et al. (2006), Sacerdote (2007)), and (iii) instrumental variables based on natural experiments (see, for example, Black et al. (2005), Currie and Moretti (2003)). We are not aware of any data on twins or adoptees that span the post-reform period in rural China. Developing instrumental variables strategy is also daunting in our case, as it is almost impossible to find credible and comparable natural experiments at two years almost one and a half decades apart.

To understand possible causal roles played by parental education and occupation, we take advantage of a battery of recent econometric advances that do not rely on the standard exclusion

[14]The distinction between genetic transmissions and environmental factors may, however, not be as sharp as it appears. It is now well-understood in Behavioral Genetics that nature and nurture interact in complex ways, and beyond a point the distinction may not be useful (Plomin et al. (2001)). For an interesting discussion on the limits to the conventional distinction between nature versus nurture, see Goldberger (1989).

restrictions required in an instrumental variables approach. First, we use the approach developed by Altonji, Elder and Taber (2005) (henceforth AET (2005)) and provide evidence on the sensitivity of the estimated intergenerational persistence with respect to different degrees of correlation between unobserved characteristics of children and parents.[15] Second, we utilize the approach developed by Klein and Vella (2009) that exploits restrictions on second moment for identification in the absence of any credible instruments (for recent applications of identification based on restrictions on second moment (heteroskedasticity), see, for example, Schroeder (2010), Emran and Hou (forthcoming), Emran and Shilpi (forthcoming)).

The central conclusion from our empirical analysis is that intergenerational occupational mobility has increased dramatically from 1988-2002 for both sons and daughters, but educational mobility has remained largely static for daughters, and it has worsened significantly for sons within a span of 14 years. The evidence from conditional correlations and multivariate OLS regressions indicates significant intergenerational occupational and educational persistence for both daughters and sons in 1988. The intergenerational persistence in educational attainment in 1988 is significantly stronger for daughters. The intergenerational persistence in the occupation choices has virtually disappeared by 2002, for both sons and daughters, but there is no evidence that educational mobility has improved. For sons, the evidence from OLS regressions indicates an increase in intergenerational persistence in educational attainment in 2002 compared to 1988.

The results from OLS regressions are somewhat qualified by the sensitivity analysis using a biprobit model following AET (2005). The sensitivity analysis shows that the observed persistence in the education of sons in 1988 can be fully explained away by very low level of positive correlations in ability across generations due to genetic or other unobserved factors (it becomes insignificant when $\rho = 0.05$).[16] In contrast, the persistence in educational attainment for daughters

[15]The AET (2005) approach can also be used to estimate lower bounds on intergenerational links that cannot be driven by unobservable common characteristics across generations. The lower bound estimate, however, requires strong observables so that it is plausible to impose the restriction that selection on observables is equal to the selection on unobservables. For a recent application of AET (2005) sensitivity analysis and lower bounds estimates to intergenerational economic mobility, see the analysis of occupational mobility in rural Nepal and Vietnam by Emran and Shilpi (2011). The AET lower bound estimates in our application, however, may not to be very informative, because the observables are not powerful enough to make the assumption of equality of selection on observables and unoservables meaningful. We, however, note that an earlier version of the paper contained the AET (2005) lower bound estimates, and the conclusions from the lower bound estimates are same as the conclusions reported here.

[16]ρ is the correlation between the error terms in parental and a child's education or occupation equations. A

in 1988 is much stronger (it remains significant even when $\rho = 0.15$). The educational persistence in the case of daughters has remained largely unchanged (marginally declined, at best) according to the AET (2005) sensitivity analysis, while the educational persistence among sons has become much stronger (it is statistically significant at the 1 percent level when $\rho = 0.20$). The AET (2005) sensitivity analysis for occupation shows that, in 1988, the intergenerational links in occupation choices remain significant when we allow for low to moderate (positive) correlation in ability across generations (they remain significant at the 5 percent level when $\rho = 0.10$). The most striking result from the sensitivity analysis is that the estimated intergenerational occupational link in 2002 is either negative or very low and statistically not significant for both sons and daughters implying equality of opportunity in occupations. The evidence from AET (2005) sensitivity analysis thus confirms the conclusion based on OLS regressions that there have been remarkable improvements in occupational mobility in rural China from 1988 to 2002. In terms of occupational mobility, the transition and economic reform thus have been nothing short of magical!

The evidence from the multivariate OLS regressions and AET sensitivity analysis is informative; it provides strong indication that although occupational mobility has improved dramatically in rural China, the educational mobility seems to have deteriorated, especially for sons. There is strong evidence that parental education affects children's education in 2002, but is there a causal effect of parental education? Unfortunately, the AET analysis does not provide us with an estimate of the causal effect of parental education on children's education. Evidence from alternative specifications of the Klein and Vella (2009) estimator shows that there is robust evidence of a causal effect of parental education on a son's years of schooling in 2002. For a daughter's education in 2002, the evidence in favor of a causal effect is much weaker; although the point estimate is large in magnitude, it is not statistically significant at the 10 percent level. In contrast, there is no evidence of a causal effect of parental education on children's education in 1988, after 10 years of market oriented economic reform that began in 1978. The persistence in educational attainment across generations (especially for sons) in 2002, after a decade and a half of the 9 years

value of $\rho = 0.05$ implies that a child with at least one parent in non-farm is 5 percentage points more likely to participate in non-farm solely due to unobserved common factors such as genetic endowment and preference.

compulsory education law seems puzzling.[17] We provide some possible explanations for the dramatic divergence between the trends in intergenerational occupational and educational mobility from 1988 to 2002.

The rest of the paper is organized as follows. Section 2 discusses the data and construction of variables. Section 3, arranged in a number of sub-sections, presents the empirical results. Section 4 provides a discussion on possible reasons behind the dramatic divergence between educational and occupational mobility. The paper ends with a summary of the findings in the conclusions.

(2) The Data

We use two rounds of Chinese Household Income Project (CHIP), 1988 to 2002 collected by a group of international and Chinese scholar with help from Chinese Academy of Social Science. We use the rural sub-sample for our analysis. We divide each year's sample into adult daughters and adult sons samples. The adult children in this paper are defined as aged 18 years or older and younger or equal to 60 years.

The CHIP surveys are repeated cross sections, and thus are suitable for an analysis of average intergenerational mobility across the spectrum of the rural society at two different time points, 1988 and 2002. As noted before, our focus is on the question whether Chinese rural society on an average has become more or less mobile over a span of one and a half decade which witnessed dramatic economic reform and spectacular growth and poverty reduction. We are comparing two snapshots of the rural society in 1988 and 2002. The data sets were collected by the same research team using exactly the same sampling methodology. Even though the questionnaire in 2002 is richer compared with the 1988 survey, it is based on the 1988 questionnaire. There is no separate parental module in the questionnaire. So the children in our sample are the children of the household head.

The 1988 data set contains 10,258 rural households in 29 provinces (or municipalities). The 2002 survey data covers 9,200 households in 22 provinces (or municipalities). Compared to the 1988 data file, it does not cover two municipalities and seven provinces (autonomous regions) , which were covered previously in 1988. The adult children samples are as follows: 3231

[17] A law mandating 9 years of compulsory education was passed in 1986. However, the implementation of the law has not been unform across provinces and counties (Tsang (2000), Hannum and Park (2007)).

(1988, daughters), 3363 (1988, sons), 2091 (2002 daughters), 3573 (2002, sons). The analysis of occupational persistence is done on the basis of these samples.

For the analysis of educational persistence, we use the sample of children born in 1967 or later. This is done to make sure that the educational attainment was not directly disrupted by Cultural Revolution, which officially lasted from 1966 to 1976. So the adult children in our samples reached the age of 8 after 1975, which means they entered primary school either at the end of the Cultural Revolution, or after the Cultural Revolution.[18] The samples for the education analysis are: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, sons). The summary statistics of the explanatory variables for the full sample are presented in appendix Table A.1.

(3) Empirical Results

(3.1) Educational Mobility

Stylized Correlations

Table 1 presents the average educational attainment and basic correlations in the data between parental education and children's education. The average years of schooling in 1988 is 5.8 years for daughters and 6.8 years for sons, which increases to 9 years for both daughters and sons in 2002. The evidence thus clearly indicates that the educational attainment in rural China has improved significantly over the 14 years, and more strikingly the initial gender gap in average education has completely disappeared by 2002.

Panel B of Table 1 shows the correlations between parents' average education (average of mother's and father's years of schooling) and a child's years of schooling. The correlation has gone down for daughters, but has increased for sons from 1988 to 2002.

Panel C of Table 1 reports the percentage of children attaining more than primary schooling conditional on parents' having more than primary schooling (6 years schooling in China). For daughters generation in 1988, the probability that a woman has more than primary education is 0.34 when none of the parents has higher than primary schooling; it increases dramatically to 0.57 when at least one parent has higher than primary schooling; the probability becomes even

[18] Sato and Li (2007a, 2007b) take a similar approach.

higher when both parents have higher than primary schooling (0.75). A daughter's probability of attaining more than primary schooling conditional on mother having more than primary is 0.70 in 1988, but it is only 0.40 when the mother has less than primary schooling. For sons in 1988, the probability conditional on father having more than primary education is 0.74, and it falls to 0.56 when the father has less than primary schooling. By 2002, the probability of having higher than primary schooling has increased dramatically for every woman irrespective of the parental educational status, but the advantage enjoyed by a daughter of more educated parents although weaker remains substantial. In 2002, a daughter is approximately 14 percentage points more likely to attain higher than primary education if at least one parent has higher than primary schooling, and by 20 percent when both parents have higher than primary schooling. The patterns for sons over time and across parent's education status are broadly similar to that observed in daughters sample.

The last panel in Table 1 reports the percentage of children attaining more than junior secondary schooling conditional on parents' education (more than primary schooling as the threshold for parents). The pattern shows that parental education affects children's probability of attaining more than junior secondary schooling in both the years, but interestingly the initial gap between sons and daughters has disappeared in 2002.

Econometric Analysis

Most of the existing econometric analysis on intergenerational educational mobility focuses on years of schooling as a measure of educational attainment. For basic regression analysis, we follow the literature and report results using years of schooling as the indicator of educational attainment. As a measure of parent's educational attainment, we use the average of father's and mother's years of schooling. We, however, use binary indicators of educational attainment later for AET (2005) sensitivity analysis, as it relies on a biprobit model.

Starting with a simple bivariate specification with no other controls, we report a series of OLS regressions with increasingly richer sets of controls. They provide suggestive evidence on the strength of selection on observables, i.e., possible roles played the observable individual, household, and county characteristics in determining the strength of the link between parental education and children's education. The control variables include observable characteristics that can proxy for ability and preference heterogeneity of parents and children. Under the assumption that selection on unobservables is similar to selection on observables, the sensitivity of the estimated intergenerational educational persistence with respect to the observables can also be informative about the possible strength of selection on unobservables (for a discussion of how selection on observables can be used as a guide to selection on unobservables, see Altonji et al (2005)). By using appropriately selected controls in the OLS regressions, we provide some suggestive evidence on the possible roles played by geography and parental wealth (income effect) in intergenerational economic mobility in rural China from 1988 to 2002.

Table 2 reports the estimated intergenerational persistence in years of schooling for alternative sets of control variables. The first four columns in Panel A of Table 2 report estimates that focus on the role of geographic location in determining the strength of intergenerational educational correlations. The second column includes a dummy for coastal province which is motivated by a large literature that shows that coastal-interior divide is important in cross-section income inequality. Interestingly, the estimated effect of parental education barely changes as we include the coastal dummy. In contrast, the estimated effect in all four cases declines significantly when we include province and county fixed effects instead of the coastal dummy (see column 3 for province fixed effect and column 4 for county fixed effect). The results in columns 1-4 thus indicate that geographic location is important for the intergenerational persistence in education, even though the coastal-interior distinction mattered little even in 1988. Also, the importance of location remains robust over a period of 14 years. The county and province fixed effects may capture, for example, availability and quality of schools across different provinces, and across counties (when county fixed effects are used), and heterogeneity in access to urban labor markets (and thus in returns to education), among other things.[19] Without these controls, it may be difficult to isolate the effects of parent's education. One might find spurious effect of parental education on children's education simply because both are driven by common factors, for example, persistent heterogeneity in labor market opportunities (nonfarm) across different geographic locations. Both parents and children may choose to become miners, because that is the only job available in a

[19]This is consistent with recent evidence that the geographic location of a rural household affects the economic outcomes significantly because of differences in access to domestic and international markets. For example, see the evidence reported by Emran and Hou (forthcoming) in the context of rural China.

mining town, and one would expect strong correlations in the educational attainment of miners working in the same mine. It may be misleading to attribute such persistence in education (or occupations) to family background or genetic inheritance.[20]

The next column reports estimated effects when a set of controls representing household wealth is included in the regression along with county fixed effects. The proxies for household wealth effect include parental occupation (farm versus non-farm), indicator of political capital (communist party membership), dummy for ethnic minority, the value of land and house, dummy for irrigation and electricity. We also include father and mother's age as part of the wealth indicators, because they capture cohort effects.[21] Once we control for the household wealth, the estimated effect of parental education declines more in all the cases, although the magnitude of the decline is not large. To understand better the relative importance of location and household wealth, column (6) reports estimates from a specification that controls for household wealth, but does not include any fixed effects (county or province). Interestingly, the results show that the effect of household wealth has a clear gender dimension. The intergenerational educational link for sons seems to be significantly sensitive to household wealth in both 1988 and 2002. In contrast, the educational correlation for daughters does not seem to depend in any significant way on household wealth in both the years. The last column in Table 2 adds a number of additional controls to the specification in column (5) including number of siblings, age of children, type of terrain in the village, old revolutionary base dummy, ethnic minority region dummy, impoverished region dummy, and suburb of big city. The results in the last column in all four cases are striking in that the addition of these variables does not affect the strength of the intergenerational educational persistence; in three out of four cases the estimated intergenerational educational link remains numerically identical, and in the only case where it changes the effect (1988 daughters), the magnitude of the change is ignorable (the estimate declines from 0.25 to 0.24). The results in Table 2 thus suggest that selection based on geographic location and household wealth are likely to be the most important for understanding intergenerational educational links in the context of

[20] One can argue that geographic location should be included as part of the family background, especially when there is no restrictions on location choices. While acknowledging this, we find it useful to separate out the role of geography from other family background variables. See below for more on this point.

[21] The conclusions about the wealth effect, however, does not depend on the precise definition of the set of wealth indicators used.

rural China. This evidence of importance of location for intergenerational educational persistence is consistent with the recent evidence that a household's location is among the most important determinants of educational attainment in China during the reform period (see, for example, Connelly and Zheng (2003)).

The potential role of geographic location in determining the strength of the educational link is interesting because migration in rural China has been restricted by Hukou system since 1951, and has been gradually relaxed, especially after 1993. Since it is likely that the parents or grand-parents chose the initial location before the imposition of Hukou system, the location can also be viewed as a summary statistic of unobserved heterogeneity (ability and preference) of parents and grand-parents.[22] Given the restrictions on geographic mobility, it is also likely that the children with rural residence live in the same village (or the same county) as their parents, and thus the county fixed effects capture common labor market opportunities due to initial economic endowment of a county (for example, mineral resources, land productivity, climatic conditions etc).[23]

The estimates of the effects of parental schooling on children's schooling reported in Table 2 tell a story which is largely consistent with the correlations reported in Table 1; the strength of the intergenerational persistence in schooling for sons is lower than that for daughters in 1988, but they have become comparable in 2002.[24]

The estimates in the last column of Table 2 (column 7) show that the intergenerational persistence in educational attainment remains both numerically substantial and statistically significant in 1988 and 2002 even after we control for a rich set of individual, parental, household, village level variables along with county fixed effects. The controls used include age of child and both mother and father, and parents' occupation, are thus likely to pick up a good measure of the genetic endowment of children. As pointed out before, the changes in the educational persistence from 1988 to 2002 are not likely to be due to changes in genetic correlations, as genetic traits do not change in any significant way in a decade and a half. We provide a more complete treatment of the possible role played by common unobserved factors such as genetic transmissions in the

[22]The adult children sample used in the paper consists of individuals born after 1967.

[23]Before 1998, a child would inherit the Hukou of his/her mother. After 1998, it can be either mother's or father's Hukou if they are different.

[24]This conclusion does not depend on the set of controls used across different columns in Table 2.

estimated intergenerational educational persistence below.

Can the Persistence in Educational Attainment Plausibly be Due to Genetic Correlations? Evidence from AET (2005) Sensitivity Analysis

As discussed in the introduction, when one finds that the intergenerational linkages in economic status are numerically and statistically significant in multivariate regression analysis, as we find above in Table 2, the estimated persistence can still be driven primarily (or solely) by unobserved heterogeneity such as transmission of genetic endowments from parents to children (both ability and preference transmission). If the observed intergenerational persistence is due to genetic correlations rather than economic environment, the scope for policy interventions to improve economic mobility in a society would be more limited.

Sensitivity Analysis

In this section, we use a bivariate Probit model to explore the question whether a small amount of selection on unobservables (such as genetic endowment) can explain away the estimated partial correlations in education of parents and children in Table 2. For the bivariate probit analysis, we use the most complete specification as in column (7) of table 2.[25] The binary indicators of educational attainment are defined as follows. For parents, the educational attainment dummy equals 1 if at least one of the parents has more than primary schooling and zero otherwise, for both 1988 and 2002. For children, the education dummy in 1988 is defined same as that for parents. But in 2002, we use middle school (9 years of schooling) as the relevant threshold for the children, as average schooling has increased substantially over the 14 years for children, and primary schooling is no longer meaningful as an indicator of educational 'attainment' (the average years of schooling is 9 years for both sons and daughters in 2002).

[25] As discussed by AET (2005), biprobit may face significant convergence problems with large number of dummies as is the case in our application, because we use county fixed effects. Since the model fails to converge with county fixed effects, we follow Emran and Shilpi (2011) and use an index of county fixed effects derived from univariate probit.

Consider the following bivariate Probit model for individual i.

$$
\begin{aligned}
E_i &= 1(\alpha E_i^p + X_i'\gamma_1 + \sum \delta_j \omega_j + \xi > 0), && (1)\\
E_i^p &= 1(X_i'\beta_1 + \sum \delta_j \omega_j + u > 0) && (2)\\
\begin{bmatrix} u \\ \xi \end{bmatrix} &\sim N\left[\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} 1 & \rho \\ \rho & 1 \end{bmatrix}\right] && (3)
\end{aligned}
$$

where E_i and E_i^p are binary measures of educational attainment of a child and the parents respectively, ω_j is the relevant fixed effect (at the county or province level) included to control for unobserved and observed community level determinants including schooling supply, labor market opportunities, agglomeration and peer effects. The error terms ξ and u represent the unobserved genetic factors for a child and her parents respectively that are relevant for their educational attainment. The correlation between ξ and u is denoted as ρ which captures the unobserved factors common across generations that can give rise to intergenerational persistence even in the absence of any causal effect of parental education.

One can argue that the above bivariate probit model is identified from nonlinearity, but such identification without any exclusion restrictions is not regarded as credible. Following AET (2005), we treat the bivariate probit model as underidentified by one parameter (ρ), and estimate the magnitudes of intergenerational educational link for different values of the correlation (ρ). The vector of explanatory variables (X) is the same as that in the regression results presented in column (7) of Table 2. Note that ρ represents only that part of genetic correlation across generations which influences the educational attainment alone, and thus is likely to be much smaller than the average genetic correlation between parents and children.[26] Also, since we include a set of control variables to proxy for ability and preference of children and parents, ρ represents only any remaining genetic influences relevant for educational attainment of both generations.

The results from sensitivity analysis are reported in Table 3. First, consider the results for 1988. The first column presents the estimated intergenerational educational persistence from

[26] The average correlation between parents and children in IQ is 0.50. But this captures both the genetic and environmental effects (Plomin et al., 2001).

univariate Probit model which assumes that $\rho = 0$; the estimates are 0.16 for daughters and 0.09 for sons, and both the estimated effects are statistically significant at the 1 percent level. Although the numerical magnitudes of the estimated effects are somewhat smaller with binary education variables, the basic conclusion remains the same as found earlier in Table 2 using years of schooling as the measure of educational attainment. The focus in the literature has been on the possibility that the estimated intergenerational persistence in column (1) of Table 3 might be driven largely by transmission of preference and ability to children from parents. Since the worry is whether the estimated effect is due to *positive selection* on unobserved genetic endowment, we implement the sensitivity analysis only for positive values of ρ. It is, however, important to appreciate that the correlation in the error terms in the triangular biprobit model represents any other factors that can affect both parental and children's education in addition to the genetic endowment emphasized in the literature.

Column 2 in Table 3 reports the estimate for $\rho = 0.05$ which implies that the child of parents with higher educational attainment (at least one parent with more than primary schooling) is five percentage points more likely to have higher educational attainment simply because of genetic transmissions or other similar positive common factors. The estimate of intergenerational persistence in education in 1988 goes down a bit in the case of daughters, from 0.16 to 0.14, but remains statistically significant at the 1 percent level. In contrast, for sons the estimate becomes almost half in magnitude and also statistically insignificant. Consistent with the a priori expectations, the estimates decline more when the value of ρ is increased to 0.10; the estimate is 0.10 (t stat=3.51) for daughters and 0.02 (t stat= 0.51) for sons. The estimated effect of parent's education on daughters educational attainment remains numerically and statistically significant even when the value of ρ goes up to 0.15. The results thus indicate that, in 1988, the observed effect of parent's education on a daughter's education is robust, and is unlikely to be driven solely by low to moderate levels of (positive) selection on unobservable genetic endowments. In contrast, the link between parents and sons is much weaker, and it can be easily explained away by very low levels of common genetic influences.

The sensitivity results for the year 2002 show a very different pattern compared to that in 1988; the intergenerational persistence in educational attainment seems to have become much stronger for sons in 2002, but it has remained largely static for daughters. The univariate probit estimates ($\rho = 0$) show that the marginal effects of parental education (having at least one parent with more than primary schooling) are similar for sons and daughters, with the effect on sons being slightly larger (0.14 for sons and 0.12 for daughters) in 2002. The second column shows that when $\rho = 0.05$, the estimates decline from 0.12 to 0.10 for daughters and from 0.14 to 0.12 for sons. The estimates for the case when $\rho = .10$ are 0.08 for daughters and 0.10 for sons, and they are statistically significant at 1 percent level (the t statistic is, however, much higher for sons). The results in Table 4 indicate that, for both sons and daughters, the estimated intergenerational link in educational attainment in the year 2002 is not wiped out by moderate levels of selection on genetically transmitted ability and preference, and the link for sons seems especially strong. The effect of parental education remains numerically substantial (0.05) and statistically significant at the 1 percent level ($t = 2.93$) for sons even with a $\rho = 0.20$, but the effect for daughters becomes close to zero and statistically insignificant.

The overall results from the biprobit sensitivity analysis thus indicate that (i) there is convincing evidence of an increase in intergenerational persistence in educational attainment of sons from 1988 to 2002, and (ii) the intergenerational persistence in education for daughters remained largely unchanged (or weakened marginally) over the same time period, and (iii) low to moderate levels of positive selection on unobservables cannot explain away the observed educational persistence in 2002, especially for sons.

Estimating the Causal Effects of Parental Education

The evidence presented so far in Tables 1-3 show interesting patterns in the intergenerational link in educational attainment in rural China. The evidence that parental education exerts strong effect on children's (especially son's) education in 2002 raises the possibility that the much-discussed increase in the cross-sectional inequality following the economic reform in China is being reinforced and reproduced by increasingly stronger intergenerational link in educational attainment. However, the evidence presented so far in Tables 1-3, while suggestive and useful, cannot provide answer to the question if the parental education has a causal effect on children's education. The AET (2005) sensitivity analysis assumes that the selection on unobservables is *positive*, and that the OLS estimates in column (7) of Table 2 are biased upward. However, the estimated effect from multivariate regressions as in column (7) of Table 2 above can be biased upward or downward because of unobserved heterogeneity, and it is not possible to pin down the direction of bias from a priori arguments. Another important issue is measurement error and the resulting attenuation bias. The importance of tackling measurement error is now well-appreciated in the literature on intergenerational income mobility. One might argue that the extent of measurement error is likely to be less in self-reported schooling data compared to measurement error in income data (for a discussion on measurement error in income and consumption data in developing countries, see Deaton (1997)). However, there is substantial evidence that measurement error in reported schooling data is a serious problem even in a highly developed country such as USA (Card (1999)). A large and mature literature on returns to education finds that the causal effects of education on earnings are, in general, higher than the corresponding OLS estimates validating the worry that the OLS estimates might suffer seriously from attenuation bias.

We implement the recently developed Kelin and Vella (2009) approach for estimation of the causal effect of parental education. The Klein and Vella (2009) approach is developed for a model with a continuous outcome and binary endogeneous treatment. We use years of schooling of children as the measure of educational attainment for the causal analysis.[27] The parental education variable remains binary with primary schooling as the threshold as in the AET (2005) sensitivity analysis above. Since we are interested in the causal effect, we carefully select the conditioning variables so that potentially endogeneous variables are excluded from the specification. The set of control variables in the results reported in Table 4 is thus different from the set used in Tables 1-3. Compared to the specification in column (7) in table 2, we exclude land and house value, parent's communist party membership and occupational status (farm vs.non-farm), dummies for irrigation and electricity, number of siblings, and dummy for female headed household. However, we note that the qualitative conclusions regarding the causal effect of parental education does not depend on the exact set of the control variables.

[27]The qualitative conclusions reached on the basis of Klein and Vella (2009) approach are robust to using alternative heteroskedastcity based estimator such as Lewbel (2011). The results from Lewbel (2011) two stage estimator are available from the authors.

Causal Effects of Parental Education: Estimates from Klein and Vella (2009) Approach

There is now a substantial literature that shows that in the absence of credible exclusion restrictions required for instrumental variables strategy, one can use heteroskedasticity for identification (see Rigobon (2003), Klein and Vella (2009, 2010),).[28] As noted by Rigobon (2003), analogous to the standard instrumental variables, heterskedasticity can be understood as a 'probabilistic shifter' of the endogeneous treatment variable which helps us trace out the causal relation between the dependent variable (children's years of schooling) and the endogeneous treatment variable (parental education). In a number of recent papers, Klein and Vella (2009, 2010) show that when at least the treatment equation in a triangular model exhibit heteroskedasticity, this effectively induces an exclusion restriction even though there is no standard exclusion restriction available. Monte Carlo evidence from a number of recent studies shows that Klein and Vella approach is effective in removing the endogeneity bias (Ebbes et al. (2009), Klein and Vella (2009, 2010), Millimet and Tchernis (2010)).

To provide intuition for the approach, we consider the following triangular model:

$$E_i^y = \alpha_0 + \alpha E_i^p + X_i\acute{}\gamma_1 + \sum_j \delta_j \omega_j + \xi_i \quad (4)$$

$$E_i^p = 1(X_i\acute{}\beta_1 + \sum \delta_j \omega_j + u_i > 0) \quad (5)$$

where E_i^y is years of schooling of child i, and E_i^p is a dummy that equals one if at least one parent has more than primary schooling. The model does not impose any exclusion restrictions on equation (4), and identification of the causal effect α is not possible if the error terms are homoskedastic.[29] Assume that the error term in the treatment equation is heteroskedastic of the following form:

$$u_i = S_u(\tilde{X}_i)\tilde{u}_i \quad (6)$$

[28] For recent applications of identification through heteroskedasticity, see for example, Farre et al. (2010), Schroeder (2010), Emran and Shilpi (forthcoming), Emran and Hou (forthcoming).

[29] As noted before, one can argue that identification in the above model can be achieved without exclusion restrictions, because the treatment equation is nonlinear. But such identification depends critically on the validity of the Normality assumption and the nonlinearity of the Normal CDF. The model is in general poorly identified. For discussions, see Kelin and Vella (2009), AET (2005).

where $\tilde{u}_i$ is a zero mean homoskedastic error, $\tilde{X}_i \subseteq X_i$ are the variables generating heteroskedasticity, and $S_u(\tilde{X}_i)$ is a positive and nonconstant function. In this case, the probability of treatment (probability of at least one parent having more than primary schooling) can be written as follows (ignoring the fixed effects):

$$\Pr\left(E_i^p = 1\right) = P\left(\frac{\tilde{X}_i}{S_u(\tilde{X}_i)}\right) \tag{7}$$

where $P(.)$ is the distribution function for $\tilde{u}_i$. With homoskedastic errors, $S(\tilde{X}_i)$ is a constant, and identification depends on possible non-linearity of $P(.)$ function such as as Normal distribution. However, such identification is based on the non-linearity in the tails of the distribution and thus relies on a small fraction of the data for identification. As such, identification based on the nonlinearity of the $P(.)$ function is in general deemed not credible. In contrast, when there is heteroskedasticity, $S(\tilde{X}_i)$ is not a constant function, and identification exploits data from the region where $P(.)$ is linear. The predicted probability of treatment from estimating equation (7) above becomes a valid instrument when there is heteroskedasticity. For the specification of the $S_u(\tilde{X}_i)$, we follow the parametric approach developed in Farre, Klein and Vella (2010) which is based on the model of heteroskedastic probit due to Harvey (1976).

$$S_u(\tilde{X}_i) = e^{\tilde{X}_i'\pi} \tag{8}$$

Since Klein and Vella (2009) rely on heteroskedasticity, credible identification hinges on two things : (1) a priori theoretical foundations for heteroskedasticity that identify the variables responsible for heteroskedasticity, and (2) formal evidence that there is substantial heteroskedasticity. We do not focus on the variables related to children, because the 'own characteristics' are expected to be the most important in explaining the variance in parent's educational attainment.[30] On theoretical grounds, the parental variables that can generate heteroskedastcity in their educational attainment (i.e., the $\tilde{X}_i$ vector) include age and locational indicators such as coastal dummy, and indicators of topography. Age captures the cohort effect and thus can represent variance in education due to changing costs of schooling over time. If schooling costs have been decreasing

[30]We, however, emphasize here that if we use the "kitchen sink" approach and include all the controls in heteroskedastic probit specification, the conclusions reached in this paper remain unaltered.

over the relevant time horizon, the variance is expected to be a negative function of age; the parents in older cohorts had less access to schools, and thus the ability cut-off for continuing education beyond primary schooling was higher, i.e, only high ability children found it worthwhile to continue schooling. This is expected to lead to lower variance in educational attainment in the older cohorts (conditional on observables included in the regression). If instead the costs of schooling had been increasing over the relevant time period, the variance in education would be a positive function of parental age. Note that the costs of education may include social and cultural costs in addition to financial costs (direct financial costs plus the opportunity costs of foregone earnings), and thus may vary across gender making the effect of age different for fathers and mothers. The coastal dummy represents variation in access to school, in the quality of schooling, and in returns to education, because the coastal areas enjoyed much better economic growth after the reform in 1978 and also government investment in public goods including schooling is more favorable. One would thus expect better access to schools in coastal areas in China, and also higher returns to education. This would reduce the ability cut-off for continuing education after primary schooling and thus increase the variance in educational attainment. The indicators of topography (dummies) can affect the variance in educational attainment because, among other things, they represent costs of access to markets and schools through their effects on the placement and route choice of transportation such as highways and rail road.[31] One might also include county (or province) fixed effects as important determinants of heteroskedasticity in parent's education. As discussed by Klein and Vella (2009, 2010), such fixed effects only account for the average differences across different groups of households, but the variance across different groups can be significantly different. In the context of rural China, the variance might reflect unobserved heterogeneity across households in a county in relevant factors such as costs of schooling and cultural differences in non-market valuation of education. Other locational indicators such as suburb of large city can also generate heteroskedasticity for similar reasons. Membership in cohesive social groups such as ethnic minority status can reduce the variance in educational attainment due to homogeneity in valuation of education and also less variance in genetic ability because of assortative matching

[31] For a discussion of the role of topography in determining access to markets in rural China, see Emran and Hou (forthcoming) and Deichman et al. (2010).

in marriage market based on ethnic identity.[32]

While the *a priori* plausibility of the variables generating heteroskedasticity is important for transparent foundation of the identification scheme, the identification ultimately depends on whether there is in fact significant heteroskedasticity in the data. To understand the nature and amount of heteroskedasticity in parental education, we report results from hetersokedastic probit model a la Harvey (1976) as discussed above. The results from the likelihood ratio test for the null of homoskedasticity are reported in Table 4. The evidence clearly shows that there is substantial heteroskedasticity in parental education, and the amount of heteroskedasticity has increased substantially from 1988 to 2002. The full results from the estimated heteroskedastic probit model are reported in appendix Table A.3. The results in appendix Table A.3 show that heteroskedasticity is driven by a few factors such as topographic indicators and parents age in 1988, but in 2002 a number of other factors are also important for heteroskedasticity including ethnic minority status and county fixed effects.[33]

An important question is if the instrument based on heteroskedasticity is strong enough to identify the causal effect. To check the strength of the instrument, we report the Kleibergen-Paap F statistics. The results in Table 4 show that in seven out of the eight cases reported, the Kleibergen-Paap F statistic is higher than the Stock et al. (2002) rule of thumb of 10 for one endogeneous variable. The only case where the instrument is relatively weak (Kleibergen-Paap F=7.43) is for daughters in 1988 when a limited set of variables is used in the heteroskedastic probit model (called specification Klein and Vella-2 in the Tables). One might wonder how much of the first stage variation is due to heteroskedasticity as opposed to the nonlinearity of the normal CDF in the probit model. This is important because identification that relies on the nonlinearity of the normal CDF in a probit model is generally deemed not credible. To check the role played by the nonlinearity of the normal CDF, we use simple probit model to generate the instrument and see how much it explains the variation in parental education, especially in comparison to the instrument generated by heteroskedasticity (i.e, from heteroskedastic probit). The Kleibergen-Paap statistics for the instrument derived from simple probit model are reported in Table 4. The

[32] The available evidence from Anthropological studies show that different ethnic groups in China value education differently. See, for example, Hansen (1999).

[33] As noted before, we use an index of county fixed effects derived from univariate probit, as heteroskedatic probit fails to converge with the county dummies. For a similar approach, see Emran and Shilpi (2011).

estimates clearly show the value of heteroskedasticity for identification; the highest F statistic for the simple probit based instrument is 5.14 and the lowest is 1.23. To appreciate the contrast better, consider the case of daughters in 2002. The Kleibergen-Paap F is 1.23 when we rely exclusively on the nonlinearity of normal CDF for identification, but with heteroskedasticity, it increases to 16.79 in Klein and Vella-2 specification and to 92.23 in the Klein and Vella-1 specification in Table 4.

The results from the Klein-Vella (2009) approach are reported in Panel A of Table 4 for the different cases. There are four sub-samples considered: (4.1) sons 1988, (4.2) daughters 1988, (4.3) sons 2002, and (4.4) daughters 2002. As benchmark, we report estimates from OLS and the normalized Inverse probability Weighted propensity score estimator (IPW) due to Hirano and Imbens (2001) and Hirnao et al. (2003). The IPW estimator relies on the assumption that there is no significant selection on unobservables (and also ignores measurement error). The Klein and Vella (2009) estimates can be compared to IPW estimates to get a sense of the role played by selection on unobservables and measurement error.

The first row in Panel A of all four cases in Table 4 shows the OLS estimate of the effects of having at least one parent with more than primary schooling on the years of schooling of children. Consistent with the results earlier, the OLS estimates indicate a significant effect of parental education on both daughters and sons, although the numerical magnitudes are somewhat different because of the change in the set of control variables. The estimates from the IPW are positive across the board, although in general smaller in magnitude (except for the daughters in 2002). An important difference with the OLS results is that the parental effect in IPW becomes insignificant at the 10 percent level for sons in 1988 (see row 2 in panel A).

Rows (3) and (4) in panel A of Table 4 report the estimated causal effect of parental education on children's years of schooling from two alternative specifications of the heteroskedastic probit model used for generating the instrument.[34] Row (3) corresponds to the case when the variables included in $\tilde{X}_i$ are all the parental characteristics including the county fixed effects. The point estimate for sons in 1988 is negative, but it is positive for daughters. However, the estimated effect in 1988, for both sons and daughters, has wide confidence interval, and the null

[34] The Klein and Vella (2009) estimates reported here are from the STATA program written by Millimet and Tchernis (2010). The confidence intervals are bootstrapped using 250 replications.

of no causal effect of parental education cannot be rejected at the 10 percent level. The 2002 estimate for sons is strikingly different; it is statistically significant at 5 percent level and also numerically substantial. According to the estimate, having at least one parent with more than primary schooling contributes two years of additional schooling for sons in 2002. In contrast, the estimated causal effect for daughters in 2002 is statistically insignificant at the 10 percent level and also numerically small. To check for the robustness of the results, we estimated the causal effect of parental education using a number of alternative specifications of the heteroskedastic probit model including the "kitchen sink" specification that includes all of the control variables in the heteroskedastic probit (i.e., including children's characteristics). As to be expected, the magnitude of the causal effect vary depending on the set of variables selected for heteroskedasticity, but the central conclusions remain intact across such alternative specifications.[35] Row (4) in panel A of Table 4 reports the estimates from one such alternative specification of the heteroskedastic probit as an example. For the alternative specification, we include the following variables in the $\tilde{X}_i$ vector: mother's age, father's age, coastal dummy and topography dummies. The estimates from this alternative heteroskedastic probit based instrument support the conclusion that there is robust evidence of a substantial causal effect of parental education only in the case of sons in 2002. In 1988, the estimates for both sons and daughters change sign compared to the earlier specification of the heteroskedastic probit. Such reversal of signs is not unexpected given the very wide confidence intervals that span both positive and negative parts of the real line. In 1988, for both sons and daughters, again the null of no parental effect cannot be rejected at the 10 percent level. In 2002, the effect of parental education on daughters is not significant at the 10 percent level. The effect is significant in the case of sons at the 5 percent level, although the numerical magnitude is somewhat smaller compared to the estimate in row (3).[36]

The results on educational mobility in Table 4 are based on the sample of children born on or after 1967. As discussed in the data section earlier, this is done to make sure that the children's education is not disrupted by the cultural revolution. However, the 1967 cut-off also implies that

[35] Since different sets of heteroskedasticity generating variables create different instruments from the heteroskedastic probit model, the estimated causal effects would naturally differ, because they provide different LATEs.

[36] If we pick the variables that are significant in the heteroskedastic probit regressions for parental education that correspond to Klein and Vella1 and Klein and Vella2 models in Table 4, and rely on them for heteroskedastcity, the estimated causal effect are consistent with the results reported in Table 4. The results are available from the authors.

the 1988 children sample consists of relatively younger cohorts (age range 18-21) compared to the 2002 sample (age range 18-35). One might wonder if the difference in the age cohorts between 1988 and 2002 samples is partly responsible for the results in Table 4. To explore this, we use a sub-sample for 2002 consisting of children of 18-21 years of age, and the estimates are consistent with the conclusions reached earlier (not reported in the Table 4). For example, the estimates from the Klein and Vella specification 1 (i.e, with all parental variables and county fixed effect used for heteroskedasticity) show that the effects on sons is 1.26 and statistically significant at the 5 percent level, but the effect on daughters is not significant at the 10 percent level (P-value 0.50).[37]

The evidence from alternative specifications of Klein and Vella (2009) estimator when combined with the evidence from the AET sensitivity analysis provide us very strong confidence in the conclusion that for sons the intergenerational persistence in education has increased significantly from 1988 to 2002, while the effect of parental education on daughters probably has not changed in any fundamental way over the same time period.

(3.2) Occupational Mobility

Stylized Correlations

Table 5 presents the basic statistics on employment status of sons and daughters over time and conditional on occupational status of parents in a given year. Panel A reports the average non-farm employment rates in our data set, and panel B shows the probability of non-farm employment for sons and daughters conditional on the occupational status of parents. Following Emran and Shilpi (2011), we define non-farm occupations as non-agricultural occupations in the rural areas (called rural non-farm activities in the literature).[38]

According to the estimates in panel A of Table 5, the (unconditional) probability of non-farm participation in 1988 is 0.22 for sons and 0.15 for daughters. The probability of non-farm participation increases dramatically in 2002, to 0.52 for sons and to 0.46 for daughters. The

[37]The heteroskedasticity based instrument is strong in the 18-21 years samples for 2002. The Kleibergen-Paap F is 44.47 in sons' sample and 36.19 in daughters' sample.

[38]The non-farm occupation includes non-farm individual enterprise owner (such as retailer, driver, etc.), employee in non-farm individual enterprise, ordinary worker, skilled worker, professional or technical worker, ordinary cadre in an enterprise, temporary or short-term contract worker.

evidence thus indicates that (i) a significant shift in favor of non-farm sector has occurred in the occupational structure in rural China over a period of 14 years, and (ii) there is a persistent gender bias against women in non-farm occupations; a daughter is about 6-7 percent less likely to participate in the non-farm occupations both in 1988 and in 2002.

The estimates in panel B of Table 5 indicate that, for both daughters and sons, the choice of nonfarm occupation depends on parent's occupational status in an important way in 1988. We consider parental occupational status to be non-farm when either or both of the parents are employed in that sector. A daughter's probability of participation in non-farm sector is 0.09 when neither of the parents work in non-farm, but it increases to 0.33 when at least one of the parents work in the non-farm sector, and to 0.71 when both parents are in non-farm. The Similar pattern also holds for sons in 1988. The influence of parental non-farm occupation has, however, become less important in 2002, the probability that a daughter of parents in agriculture has non-farm occupation has increased from 0.09 in 1988 to 0.43, the probability for a daughter with both parents in non-farm has remained virtually unchanged (0.73 in 2002 and 0.71 in 1988). A similar pattern holds for sons. The evidence thus suggest that the advantage enjoyed by the children of non-farm parents has weakened dramatically in a period of 14 years from 1988 to 2002.

Econometric Analysis

Following the approach adopted for the educational mobility results in Table 2, we report the results on occupational persistence sequentially, starting from a simple Probit regression we introduce an array of control variables in subsequent steps. This helps demonstrate the robustness (or non-robustness) of intergenerational linkages in non-farm participation.The results from a series of Probit regressions are reported in Table 6.[39] The parental occupational status is considered to be non-farm when at least one parent is employed in that sector. The reported estimates are the marginal effects of at least one parent in non-farm on the probability of a child's non-farm participation.

The pattern of sensitivity of the estimated intergenerational persistence coefficient across different sets of control variables in Table 6 is broadly similar across gender, and interestingly, mimics reasonably well the pattern found earlier in the case of intergenerational educational

[39]The results from linear probability model are very similar and thus omitted for the sake of brevity.

persistence, especially in 1988. The estimated marginal effect of having at least one parent in nonfarm sector in 1988 is about 0.25 for both sons and daughters when no controls are included in the regressions. The estimates change only marginally when we include a coastal dummy, thus confirming the finding in the case of education earlier that the coast versus interior distinction is not very pertinent for intergenerational persistence in economic status in rural China. The third and fourth columns report results that include province and county fixed effects respectively without any additional controls. The estimated effect of parental education in 1988 goes down for both daughters and sons with province fixed effects, the magnitude of the decline being similar across gender. However, when we include county fixed effects instead of province fixed effects, the decline in the marginal effects of parental non-farm participation is much larger in the case of daughters (column 4). The cross-county differences in 1988 thus seem to be especially important in determining the strength of intergenerational occupational persistence for daughters. The specification in column 5 includes a set of indicators of household wealth including land in addition to the county fixed effects employed in the previous column. The estimated intergenerational occupational linkage in 1988 for both daughters and sons become even smaller, thus confirming the importance of wealth effect. However, similar to the results for education, the effects of wealth variables depend on the locational fixed effects. If we include the wealth indicators without any fixed effects (column 6), they do not affect the magnitude of intergenerational linkage in any appreciable manner (compare columns 4 and 6). The last column in Table 6 shows the estimates when we include additional variables that capture geographic and household level heterogeneity; the estimated effect of parental non-farm participation in 1988 goes down a bit more for daughters, but remains unchanged for sons. The estimates in Table 6 thus indicate that parental occupational choices had had significant influence on children's occupation choices in 1988; even after controlling for a rich set of individual, parental, household, village characteristics along with county fixed effects, the estimates in column 6 imply that a son or daughter is 12 percentage point more likely to choose nonfarm occupation when at least one parent is in non-farm sector.

But the most striking evidence in Table 6 relates to the evolution of intergenerational persistence in occupation choices over 14 years from 1988 to 2002; the intergenerational link between parents and children in non-farm participation has disappeared in 2002, both for daughters and sons. For daughters, the coefficient on the dummy for 'at least one parent in non-farm' is positive but small and not significant at the 10 percent level when we include only province fixed effects, and it turns negative and insignificant once county fixed effects are used. For sons, the strength of the intergenerational link in 2002 is a bit stronger, but it also does not survive when we include county fixed effects. The evidence thus shows that occupational mobility has dramatically improved in rural China; by 2002, a farmer's children are no longer doomed to be farmers themselves. In terms of occupational mobility, the transition to more market oriented open economy has been nothing less than magical in rural China.

Unobserved Heterogeneity and Occupational Linkages

The results from Probit regressions in Table 6 provides convincing evidence that occupational mobility has improved in rural China from 1988 to 2002. The observed change in the intergenerational occupational linkages cannot be due to genetic transmissions from parents to children, as the genetic traits do not change in a decade and a half. The evidence also indicates that the link between parental occupation choices and children's occupation remained strong in 1988 even after controlling for a rich set of observable characteristics. However, the observed persistence in column (7) of Table 6 in 1988 can still be driven primarily by unobserved heterogeneity, such as positive selection on common genetic endowments and preferences. Since the treatment and the outcome variables are both binary, it is natural to take advantage of the AET (2005) sensitivity analysis to explore if low to moderate positive selection on common unobserved ability and preference can explain away the observed occupational linkages in 1988. For the sake of completeness we also report the sensitivity results for 2002. Note that we do not use the Kelin and Vella (2009) approach to estimate the causal effects in the case of occupation as it requires a continuous outcome variable. If we ignore the binary nature of the dependent variable and implement Klein and Vella (2009) approach, the conclusions regarding the evolution in intergenerational occupational linkages reported in this paper remain intact.

The results from the biprobit sensitivity analysis for alternative values of ρ are presented in Table 7. The first column reports the estimate from univariate probit that assumes that $\rho = 0$ which implies in the present context that the genetic transmissions of ability and preference from parents to children can be treated as insignificant. The second column reports the estimates for $\rho = 0.05$ implying that children with at least one parent in the non-farm occupation would be 5 percentage points more likely to choose non-farm occupation simply because of genetic transmissions or other similar unobserved factors. The estimated effect in 1988 declines from 0.12 to 0.09 for daughters and from 0.12 to 0.10 for sons, but it remains statistically significant in both cases at the 1 percent level. The effect of parental non-farm participation on both sons and daughters remain statistically significant at the 5 percent level even when ρ is increased to 0.10. The evidence thus indicates that the occupational linkages in column (7) of Table 6 in 1988 are robust to allowing for moderate positive correlations between the parents' and children's unobserved characteristics. Interestingly, the occupational persistence appears to be more robust in the case of daughters, it remains significant at the 10 percent level when the value of ρ is as high as 0.15, but the effect in case of sons becomes insignificant at the 10 percent level.

The sensitivity analysis for 2002 confirms the conclusion that in 2002, there is little or no persistence in occupation choices across generations in rural China, once the effects of location is taken into account.[40]

(4) Towards an Understanding of the Divergence Between Educational and Occupational Mobility

Probably the most striking finding from our empirical analysis is that intergenerational occupational persistence in rural China has effectively vanished in a span of 14 years. Another interesting aspect of the dramatic improvements in occupational mobility is that its benefits are distributed equally across gender. Such a dramatic improvement in occupational mobility for both sons and daughters in a span of a decade and a half is no less than magical. But the not so magical part of the transition to a more market oriented economy in rural China is that the educational mobility has not witnessed any noticeable improvement over the same period of time. The daughters fared better in terms of educational mobility; but even they did not experience any significant improvement from 1988 to 2002. For sons, the evidence in fact suggests that, if anything, the intergenerational persistence in educational attainment has become much stronger, implying significantly lower mobility in 2002. While a complete treatment of possible reasons behind this divergence between occupational and educational mobility is beyond the scope of the

[40]The high occupational mobility in rural China in 2002 is similar to the evidence found in Emran and Shilpi (2011) for rural Vietnam using LSMS 1992/93 data.

present paper, we put forth a number of possible explanations behind the observed pattern in intergenerational persistence which can be explored in depth in future research.

One might expect that occupational mobility and educational mobility should go hand in hand in rural areas, especially given the evidence from recent research that probability of rural non-farm employment increases with education (Yang (1997), deBrauw and Rozelle (2008)). However, note that most of the rural non-farm occupations do not require more than primary schooling, and the proportion of children with at least primary education has increased substantially; it was 97 percent for daughters and 100 percent for sons in 2002. So by 2002 education was not a constraining factor for most of the rural children for participation in non-farm activities.[41] The impressive productivity growth in agriculture following the implementation of household responsibility system meant that the rural households did not need the children to work on the farm to produce enough food; the children could explore alternative occupations without facing the prospect of quasi starvation. On the demand side, the spectacular growth of the non-farm sector fueled by growth in the urban income and also expanding export market for the products of TVEs ensured that there was enough demand for rural non-farm products.

The lack of improvements (worsening in the case of sons) in educational mobility can be traced to a host of factors including increased direct cost of education, and higher opportunity cost of continuing in the school. Although China adopted a legislation in 1986 for compulsory 9 years of education, its implementation has not been uniform, the rural areas have in general lagged behind (Behrman et al. (2008), Ma and Ding (2008), Tsang (1994)). The education reform focused on quality of education and shut down some low quality schools. Fiscal decentralization tightened the link between local economic conditions and educational opportunities in a village (Hannum et al. (2007b), Tsang (2002)). Even though the central government provided transfers to the poor areas, the schools in poor areas were forced to cover costs by charging various fees to their families. Tuition and fees increased from 4.42 percent of household expenditure in 1991 to 18.59 percent by 2004 (Behrman et al. (2008)).[42] The increased monetary costs of education naturally increased

[41] This can also be demonstrated by looking at the probability of a child attaining primary schooling conditional on primary schooling of parents in 2002. By 2002, for attaining primary schooling, the parental education mattered little.

[42] For detailed and in-depth evidence on the importance of fees and related costs in schools of rural China, see Hannum (2008).

In a World bank report, Piazza and Liang (1998) conclude that "despite the extraordinary success in basic the persistence in educational attainment across generations; only the relatively rich in rural areas could afford education. After the economic reform in 1978, the relatively educated parents could take full advantage of the new opportunities both in agriculture and non-farm sector, and thus they reaped high income. Yang (2004) shows that the more educated people were able to allocate their resources more efficiently under the household responsibility system (see also Li and Zhang (1998)).[43] This higher income allowed them to invest in children's education in the face of increasing private cost of schooling. The disadvantage faced by the children of less educated (and thus poorer) parents was reinforced by the rising returns to working in the rural non-farm sector (or urban migration); the opportunity costs of not working along with higher monetary costs of education made it difficult for them to continue schooling beyond a low threshold (primary schooling for example). Another factor that may have contributed to the educational immobility is the increasing returns to education in rural China following the economic reform (Behrman et al. (2008)). As shown by Solon (2004) higher returns to education is expected to increase the persistence in intergenerational economic status.

(5) Conclusions

Using two rounds of CHIP household survey data for 1988 and 2002, this paper provides evidence on the evolution of intergenerational economic mobility in rural China. It aims to answer the following question: if we compare two snapshots of the cross-section of households in rural China, has the rural society become more or less mobile from 1988 to 2002? We use educational attainment and occupational choices as two salient indicators of economic status in rural China.

A central issue in the literature on intergenerational persistence in economic status has been the possibility that the estimated partial correlation from regression analysis may be driven primarily by unobserved heterogeneity such as genetic transmissions from parents to children. Since our focus is on the changes in the intergenerational linkages over a period of a decade and a half, the estimated link between parents and children in education and occupations from multivariate

education in China, many poor were not reached by the government efforts...in the poorer half of the townships of 35 counties supported by a World Bank projects, average enrollment was at least 10 % points lower than the national average for the same age group...".

[43] The returns to education was, in contrast, very low under collective agriculture before 1978.

regressions are likely to identify the direction of change in the intergenerational linkages reasonably well. To investigate the possible role played by positive selection on unobserved characteristics in determining the strength of intergenerational persistence in a given year, we take advantage of the sensitivity analysis developed by Altonji, Elder and Taber (2005). To estimate the causal effect of parental education on children's years of schooling, we use Klein and Vella (2009) approach that exploits heteroskedasticity for identification. The results from the empirical analysis show that the intergenerational occupational mobility has increased dramatically in rural China from 1988 to 2002. Although in 1988, parents seem to exert considerable influence on children's occupation choices, the effect of parental non-farm participation on children's non-farm choice is effectively zero in 2002, for both daughters and sons. In contrast, the intergenerational educational mobility has remained largely static for daughters, while it has become significantly less mobile over time for sons. We provide possible explanations for such divergence in occupational and education mobility in rural China in the post-reform period.

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Table 1: Descriptive statistics on education of adult children and their parents

Panel 1A: Average years of schooling

	Daughters' Sample			Sons' Sample			Full sample		
Year	**Daughter**	**Mother**	**Father**	**Son**	**Mother**	**Father**	**children**	**Mother**	**Father**
1988	5.8	2.4	5.2	6.8	2.3	5.0	6.3	2.4	5.1
2002	9.0	5.3	7.1	9.0	4.9	6.9	9.0	5.0	7.0

NOTES

1. children were born in 1967 or later;
2. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, daughters).

Panel 1B: Correlation Between Parents' Average Education and Children's Education

	Daughters' Sample	Sons' Sample
1988	0.37	0.24
2002	0.32	0.33

NOTES

1. Children were born in 1967 or later;
2. Children's education is measured by years of schooling;
3. Parents' average education is the average of mother's and father's years of schooling;
4. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, daughters).

Panel 1C: Percentage of Children with More than Primary Schooling (6 Years of Schooling) Conditional on Parent's Education

Adult children	Neither parent's education is more than primary school	At least one parent's education is more than primary school	Both parents' education are more than primary school
Daughters			
1988	34.3	56.6	74.8
2002	77.7	91.7	97.8
Sons			
1988	55.1	73.1	84.9
2002	81.4	94.2	97.6

NOTES

1. Children were born in 1967 or later;

2. Adult children's education dummy: 1=more than primary school;

3. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, daughters).

Panel 1D: Percentage of Children with More than Junior Secondary Schooling (9 Years of Schooling) Conditional on Parent's Education

Adult children	Neither parents' education is more than primary school	At least one parent's education is more than primary school	Both parents' education are more than primary school
Daughters			
1988	3.4	11.1	25.1
2002	18.4	34.7	43.5
Sons			
1988	5.3	14.7	29.4
2002	16.9	35.2	43.2

NOTES

1. Children were born in 1967 or later;

2. Adult children's education dummy: 1=more than junior secondary school;

3. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, daughters).

Table 2: Impact of Parents' Average Years of Schooling on Children's Years of Schooling

Adult Children	**Parents' edu**	**Parents'edu + coastal dummy**	**Parents' edu +province fixed effects**	**Parents' edu + county fixed effects**	**Parents' edu + county fixed effect + HH wealth**	**Parents' edu + HH wealth**	**Parents' edu + county fixed effect + HH wealth + other controls**
	(1)	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
1988--daughters							
Marginal Effect	**0.37**	**0.37**	**0.33**	**0.28**	**0.25**	**0.34**	**0.24**
t statistics	18.35	18.32	14.02	10.49	7.64	13.49	7.37
1988--sons							
Marginal Effect	**0.23**	**0.22**	**0.21**	**0.17**	**0.14**	**0.18**	**0.14**
T statistics	11.92	11.71	9.61	6.97	4.85	6.42	4.81
2002--daughters							
Marginal Effect	**0.32**	**0.32**	**0.28**	**0.23**	**0.20**	**0.30**	**0.20**
T statistics	16.09	15.83	13.44	11.01	8.83	13.89	8.85
2002--sons							
Marginal Effect	**0.33**	**0.32**	**0.28**	**0.27**	**0.22**	**0.25**	**0.22**
T statistics	20.86	20.38	17.41	15.64	11.43	14.21	11.37

NOTES

1. Adult children were born in 1967 or later;
2. Parents' average years of schooling is the average of mother's and father's years of schooling;
3. "HH wealth" stands for indicators of household wealth: includes mother's age, father's age, parents' occupation dummy (non-farm=1), communist party member dummy, ethnic minority dummy, number of grandparents alive, female household head dummy, irrigation dummy, land, house value and electricity dummy;
4. "Other controls " include topography and coastal location dummy; the number of adult male and female children in the household, old revolutionary base dummy, ethnic minority region, impoverished region dummy, and suburb of big city and adult child's age and age squared;
5. All standard errors are corrected for heteroscedasticity.

Table 3: Sensitivity Analysis of the Impact of Parent's Education on Children's Education [AET (2005)]

Sample	Year	Correlation Between the Error Terms in the Triangular Model					
		$\rho = 0$	$\rho = 0.05$	$\rho = 0.10$	$\rho = 0.15$	$\rho = 0.20$	$\rho = 0.25$
Daughters	1988	**0.16**	**0.14**	**0.11**	**0.08**	0.04	0.001
		(5.89)***	(4.55)***	(3.51)***	(2.46)**	(1.40)	(0.32)
	2002	**0.12**	**0.10**	**0.08**	**0.05**	0.02	-0.002
		(5.21)***	(4.61)***	(3.45)***	(2.28)**	(1.09)	(-0.13)
Sons	1988	**0.09**	0.05	0.02	-0.02	-0.05	-0.08
		(3.44)***	(1.49)	(0.51)	(-0.48)	(-1.48)	(-2.49)
	2002	**0.14**	**0.12**	**0.10**	**0.08**	**0.05**	0.02
		(8.33)***	(7.61)***	(6.07)***	(4.51)***	(2.93)***	(1.31)

NOTES

1. t statistics in parentheses;
2. Adult children's education dummy: 1=more than primary school for the 1988 sample;
3. Adult children's education dummy: 1=more than middle school for the 2002 sample;
4. Parent's education dummy, 1= if at least one parent's education is more than primary school;
5. "*" significant at 10 percent; "**" significant a 5 percent; "***" significant at 1 percent;
6. The estimation uses STATA routine of AET (2005);
7. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, sons).

Table 4: Causal effects of Parent's Education on Children's Education: Klein and Vella (2009) Approach

	Daughters			Sons		
	Estimate	90 Percent Conf. Interval		Estimate	90 Percent Conf. Interval	
		lower bound	upper bound		lower bound	upper bound
1988						
OLS	**1.029**	**0.674**	**1.384**	**0.524**	**0.104**	**0.809**
IPW	**0.708**	**0.25**	**1.15**	**0.111**	**-0.464**	**0.655**
Klein and Vella-1	**3.443**	**-40.656**	**25.045**	**-0.417**	**-10.041**	**20.266**
Likelihood-ratio test of heteroskedasticity		21.17			18.72	
Kleibergen-Paap F statistic		28.21			40.16	
Klein and Vella-2	**-0.998**	**-13.38**	**10.91**	**0.937**	**-15.65**	**13.407**
Likelihood-ratio test of heteroskedasticity		10.94			4.03	
Kleibergen-Paap F statistic		7.43			14.25	
Kleibergen-Paap F stat. for the instrument based on simple Probit		2.76			5.13	
2002						
OLS	**0.793**	**0.632**	**0.946**	**0.924**	**0.791**	**1.038**
IPW	**0.811**	**0.559**	**0.955**	**0.879**	**0.756**	**1.001**
Klein and Vella-1	**0.613**	**-1.16**	**5.425**	**2.067**	**0.966**	**2.826**
Likelihood-ratio test of heteroskedasticity		58.84			60.63	
Kleibergen-Paap F statistic		92.23			72.95	
Klein and Vella-2	**1.311**	**-1.089**	**3.491**	**1.583**	**0.083**	**3.341**
Likelihood-ratio test of heteroskedasticity		15.01			31.73	
Kleibergen-Paap F statistic		16.79			38.82	
Kleibergen-Paap F stat. for the instrument based on simple Probit		1.23			4.62	

NOTES
1. All regressions include county fixed effects, 90% confidence interval is boostrapped using 250 replications;
2. Controls include individual's age, age squared, mother's age and father's age, parents' ethnic minority, location dummies (topography, old revolutionary region, ethnic minority region, impoverished region, suburb of big city, and coastal region dummies);
3. IPW = normalized inverse probability weighted propensity score estimator of Hirano and Imbens (2001);
4. "Klein and Vella-1 Estimate" uses parents characteristics, location dummies and index of county fixed effect for heteroscadesicity; "Klein and Vella-2 Estimate" uses father age, mother's age, coastal dummy, and topography for heteroscadesicity;
5. Parents' education level is a binary measure, it is equal to 1 if at least one of parents has more than primary schooling; children's education is measured by years of schooling;
6. Number of Observations: 2057 (1988, daughters), 1884 (1988, sons), 2264 (2002, daughters), 3616 (2002, sons).

Table 5: Descriptive Statistics for Occupation

Panel 5A: Probability of Non-farm Occupation for Adult Children and their Parents

	Daughters' Sample			Sons' Sample			Full Sample		
Year	**Daughter**	**Mother**	**Father**	**Son**	**Mother**	**Father**	**children**	**Mother**	**Father**
1988	0.15	0.04	0.21	0.22	0.04	0.20	0.18	0.04	0.21
2002	0.47	0.08	0.34	0.52	0.08	0.29	0.50	0.08	0.30

NOTES

1. Non-farm occupation dummy=1 if involved in non-farming for 3 months or more in the survey year;
2. Number of Observations: 3231 (1988, daughters), 3363 (1988, sons), 2091 (2002 daughters), 3573 (2002, sons).

Panel 5B: Percentage of Children in Non-farm conditional on Parent's Occupation

Adult children	Neither parent is in Non-farm Occupation	At Least One Parent is in Non-farm Occupation	Both Parents are in Non-farm Occupation
Daughters			
1988	9.4	33.2	71.1
2002	43.0	52.3	72.8
Sons			
1988	16.4	41.7	73.9
2002	47.4	60.9	75.7

NOTE:

1. Non-farm occupation dummy=1 if involved in non-farm for 3 months or more in the survey year;
2. Number of Observations: 3231 (1988, daughters), 3363 (1988, sons), 2091 (2002 daughters), 3573 (2002, sons).

Table 6: The impact of parents' Non-farm participation on Children's Non-farm participation (Probit)

Adult Children	Parents' Occ.	Parents' Occ. + Coastal Dummy	Parents' Occ. + Province Fixed Effect	Parents' Occ. + County Fixed Effect	Parents' Occ. +County Fixed Effect + HH Wealth	Parents' Occ. + HH Wealth	Parents' Occ. + County Fixed Effect + HH Wealth + Other Controls
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1988—daughters							
Marginal Effect	**0.24**	**0.22**	**0.18**	**0.18**	**0.15**	**0.17**	**0.12**
T statistics	15.88	14.78	12.26	7.54	5.01	9.73	3.73
1988—sons							
Marginal Effect	**0.25**	**0.24**	**0.21**	**0.16**	**0.12**	**0.18**	**0.12**
T statistics	13.71	12.92	11.57	6.66	3.47	7.72	3.44
2002—daughters							
Marginal Effect	**0.09**	**0.08**	0.04	-0.04	**-**	**-**	**-**
T statistics	4.06	3.35	1.59	-1.61	-	-	-
2002—sons							
Marginal Effect	**0.14**	**0.11**	**0.07**	0.05	0.04	0.04	0.04
T statistics	7.45	6.19	3.59	1.59	1.52	1.44	1.42

NOTES

1. Dependent variable is adult children's occupation dummy, 1=Non-farm; parents' occupation dummy, 1=at least one parent is in non-farm;
2. "HH wealth" stands for indicators of household wealth: includes parents' communist party member dummy, ethnic minority dummy, female household head dummy, mother's age, father's age, irrigation dummy, land, house value and electricity dummy;
4. "Other controls " include topography and coastal location dummy; labor constraint; mother's and father's years of schooling, old revolutionary base dummy, ethnic minority region, impoverished region dummy, and suburb of big city and adult child's age and age squared, years of schooling;
5. All standard errors are corrected for heteroscedasticity;
6. Number of Observations: 3231 (1988, daughters), 3363 (1988, sons), 2091 (2002 daughters), 3573 (2002, sons).

Table 7: Sensitivity Analysis of the Impact of Parents' occupation [AET (2005)]

Sample	Year	Correlation of the Error Terms in the Triangular Model					
		$\rho = 0$	$\rho = 0.05$	$\rho = 0.10$	$\rho = 0.15$	$\rho = 0.20$	$\rho = 0.25$
Daughters	1988	**0.12**	**0.09**	**0.05**	**0.03**	0.02	0.00
		(3.73)***	(3.11)***	(2.42)**	(1.69)*	(0.82)	(0.03)
	2002	-0.10	_	_	_	_	_
		(-3.18)	_	_	_	_	_
Sons	1988	**0.12**	**0.10**	**0.07**	0.04	0.01	-0.02
		(3.49)***	(3.17)***	(2.24)**	(1.31)	(-0.34)	(-0.62)
	2002	0.04	0.03	0.001	-0.03	_	_
		(1.42)	(1.31)	(0.18)	(-1.32)	_	_

NOTES

1. t statistics in parentheses;
2. Adult children's occupation dummy: 1=participation in Non-farm;
3. Parent's occupation dummy, 1=at least one of parents is in Non-farm;
4. "*" significant at 10 percent; "**" significant a 5 percent; "***" significant at 1 percent;
5. The estimation uses STATA routine of AET (2005);
6. The samples are: 3231 (1988, daughters), 3363 (1988, sons), 2091 (2002 daughters), 3573 (2002, sons).

APPENDIX

Appendix A1: Descriptive statistics of variables in education estimation

Category	Description	1988 Mean	1988 Min	1988 Max	1988 Std. dev.	2002 Mean	2002 Min	2002 Max	2002 Std. dev.
Dependent variable	Dummy, 1=if the adult child's education is more than primary school	0.52	0	1	0.49				
	Dummy, 1=if the adult child's education is more than junior middle school					0.29	0	1	0.46
	Daughters' education, years of schooling	5.8	0	14	2.2	9.0	0	16	2.4
	Sons' education, years of schooling	6.8	0	14	2.7	9.0	0	16	2.4
Parents' education	Dummy, 1=either parents' education more than primary school	0.35	0	1	0.48	0.66	0	1	0.47
	Average of mother's and father's years of schooling	3.7	0	14	2.6	6.0	0	15	2.3
Individual characteristics	Age	19.43	18	21	1.1	22.63	18	35	3.9
	Age squared	379	324	441	43	527	324	1225	193
Parents' characteristics	Mother's age	46.53	35	81	5.88	48.61	35	76	5.84
	Father's age	49.16	35	74	6.28	50.96	35	82	6.33
	Dummy, 1=if either parents in non-farming	0.23	0	1	0.42	0.37	0	1	0.48
	Dummy, 1=if either parents is Communist Party member or cadre	0.2	0	1	0.4	0.41	0	1	0.49
	Dummy, 1=if either parents is ethnic minority	0.08	0	1	0.26	0.14	0	1	0.35
Household characteristics	Number of adult male children	1.79	0	6	0.98	1.32	0	1	0.82
	Number of adult female children	1.77	0	7	1.15	0.93	0	1	0.86
	Dummy, 1=if female household head	0.03	0	1	0.16	0.03	0	1	0.18
	Dummy, 1=if irrigation available to the household	0.58	0	1	0.49	0.31	0	1	0.46
	Land, unit is *mu*	14.27	0.3	956	39.9	5.91	0	180	7.09
	House value, unit is current *yuan*	5307	3	80000	5730	26596	450	360000	29684
	Dummy, 1=if electricity available to household	0.87	0	1	0.34	0.99	0	1	0.06
Village characteristics	Dummy,1= if topography is terrain (base is flat)	0.31	0	1	0.46	0.33	0	1	0.47
	Dummy, 1=if topography is mountain (base is flat)	0.19	0	1	0.39	0.19	0	1	0.39
	Dummy, 1=if old revolutionary base	0.15	0	1	0.35	0.22	0	1	0.42
	Dummy, 1= if ethnic minority region	0.08	0	1	0.28	0.13	0	1	0.34
	Dummy, 1= if impoverished region	0.19	0	1	0.39	0.28	0	1	0.45
	Dummy, 1= if suburb of big city	0.01	0	1	0.11	0.08	0	1	0.27
	Dummy, 1= if along coastal line	0.24	0	1	0.43	0.25	0	1	0.43

NOTE: Adult children born on or after 1967

Appendix A2: Descriptive statistics of variables in occupation estimation

Category	Description	1988				2002			
		Mean	Min	Max	Std. dev.	Mean	Min	Max	Std. dev.
Dependent variable	Dummy, 1 if the adult child participate in non-farming	0.19	0	1	0.39	0.50	0	1	0.49
Parents' occupation	Dummy, 1=if either parents is in non-farming	0.22	0	1	0.41	0.32	0	1	0.46
Individual characteristics	Age	21.77	18	50	3.51	23.56	18	52	4.61
	Age squared	486	324	2500	177	576	324	2704	249
	Years of schooling	6.44	2	14	2.59	8.54	0	16	2.17
Parents' characteristics	Mother's age	48.3	35	81	6.23	49.56	35	82	6.21
	Mother's years of schooling	2.18	0	14	3.04	4.79	0	15	2.83
	Father's age	50.81	35	80	6.49	51.91	35	87	6.59
	Father's years of schooling	4.87	0	14	3.12	6.73	0	15	2.56
	Dummy, 1=if either parents is ethnic minority	0.08	0	1	0.26	0.15	0	1	0.35
	Dummy, 1 if either parents is Communist Party member	0.19	0	1	0.39	0.21	0	1	0.41
	Dummy, 1 if either parents is cadre	0.05	0	1	0.23	0.33	0	1	0.47
Household characteristics	Dummy, 1= if female household head	0.03	0	1	0.16	0.03	0	1	0.15
	Number of adults per mu of land	0.63	0.01	13.3	0.66	1.01	0.02	62.5	2.06
	Proportion of female adults to the number of total adults	0.5	0	1	0.16	0.49	0	1	0.16
	Ratio of young children (younger than 7 years) to total adult women	0.16	0	3	0.31	0.11	0	2	0.11
	Ratio of children (between 8 and 17 years) to the household size	0.17	0	0.625	0.14	0.09	0	0.57	0.13
	Proportion of student in the household	0.02	0	0.57	0.06	0.08	0	0.5	0.12
	Share of disabled or retired in household	0.002	0	0.667	0.03	0.01	0	0.66	0.05
	Dummy, 1=if irrigation available to the household	0.6	0	1	0.48	0.3	0	1	0.46
	Land, unit is *mu*	14.31	0.3	959	42.41	5.99	0	180	7.14
	House value, unit is current *yuan*	5584	3	80000	6034	25477	450	300000	28425
	Dummy, 1=access to electricity	0.86	0	1	0.34	0.99	0	1	0.06
	Productive assets (current *yuan*)	1448	20	35250	1925	5244	0	793460	17872
	Savings, unit is current *yuan*	202	0	11450	788	2088	0	180000	8378
	Debt, unit is current *yuan*	243	0	36000	1014	872	0	200000	5200
Village characteristics	Dummy,1 if topography is terrain (base is flat)	0.3	0	1	0.46	0.33	0	1	0.47
	Dummy, 1 if topography is mountain (base is flat)	0.18	0	1	0.38	0.21	0	1	0.41
	Dummy, 1 if old revolutionary base	0.14	0	1	0.35	0.21	0	1	0.41
	Dummy, 1 if ethnic minority region	0.08	0	1	0.27	0.14	0	1	0.35
	Dummy, 1 if impoverished region	0.19	0	1	0.39	0.31	0	1	0.46
	Dummy, if suburb of big city	0.01	0	1	0.11	0.07	0	1	0.25
	Dummy, if along coastal line	0.27	0	1	0.44	0.23	0	1	0.42

Appendix A3: Determinants of Heteroskedasticity (from Heteroskedastic Probit)

	Daughters				Sons			
	Klein and Vella - 1		Klein and Vella - 2		Klein and Vella - 1		Klein and Vella - 2	
	Coef.	Z stat.	Coef.	Z stat.	Coef.	Z stat.	Coef.	Z stat.
1988								
Mother age	-0.03	-1.37	-0.05	(-2.12)**	-0.02	-0.88	-0.03	(-1.69)*
Father age	0.00	0.03	-0.02	-0.92	-0.00	-0.03	0.01	0.65
Coast dummy	-0.02	-0.08	-0.25	-1.07	-0.34	-1.58	-0.19	0.91
Topography (terrain)	0.04	0.11	9.88	1.34	-0.27	-1.41	-0.08	-0.31
Topography (mountain)	-0.59	(-2.16)**	0.64	1.08	-0.41	(-1.61)*	0.13	0.35
Parents' ethnic minority	0.18	0.41			-1.67	(-1.79)*		
Old revolutionary base	-0.02	-0.06			-0.09	-0.35		
Ethnic minority region	-0.00	-0.02			1.42	1.47		
Impoverished region	7.81	0.89			0.26	0.73		
Suburb of big city	-0.39	-0.65			-1.67	(-1.84)*		
County index	-0.02	-0.13			-0.61	(-4.44)***		
2002								
Mother age	-0.04	-1.61	-0.01	-0.24	-0.05	(-2.69)***	-0.04	(-2.14)***
Father age	0.17	6.31***	0.03	1.86*	0.06	3.78***	0.04	2.95***
Coast dummy	0.74	2.76***	0.28	0.79	0.63	3.31***	0.32	1.88*
Topography (terrain)	0.29	2.17**	0.11	0.59	-0.22	(-1.94)*	-0.27	(-2.51)**
Topography (mountain)	2.09	5.39***	0.37	1.28	0.44	2.76***	0.23	1.57
Parents' ethnic minority	-0.51	-1.39			-0.43	(-1.63)*		
Old revolutionary base	-2.08	(-4.31)***			0.44	2.76***		
Ethnic minority region	-1.29	(-2.41)**			0.13	0.47		
Impoverished region	-0.11	-0.68			0.01	0.11		
Suburb of big city	0.41	1.04			-0.12	-0.69		
County index	-0.27	(-2.33)**			-0.14	-1.21		

NOTES:

1. "Klein and Vella-1 Estimate" uses parents characteristics, location dummies and index of county fixed effect for heteroscadesicity; "Klein and Vella-2 Estimate" uses father age, mother's age, coastal dummy, and topography for heteroscadesicity;
2. Parents' education level is a binary measure, it is equal to 1 if at least one of parents has more than primary schooling; children's education is measured by years of schooling;
3. "***", "**", "*" stands for significance of 1%, 5%, and 10%, respectively.