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INTERGENERATIONAL INCOME MOBILITY IN URBAN CHINA

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This paper estimates the intergenerational income elasticity for urban China, paying careful attention to the potential biases induced by income fluctuations and life cycle effects. Our preferred estimate indicates that the intergenerational income elasticity for father–son is 0.63. This suggests that while China has experienced rapid growth of absolute incomes, the relative position of children in the distribution is largely related to their parents' incomes. By investigating possible causal channels, we find that parental education plays one of the most important roles in transmitting economic status from parents to children.

JEL Codes: D10, D31

Keywords: intergenerational mobility, political party membership, transgenerational persistence

1. INTRODUCTION

Economists have long been interested in the issue of intergenerational mobility. Estimating the relationship between the permanent incomes of parents and children is a critical component of a society's income dynamics. A growing body of research has demonstrated large and systematic differences across nations, with parental income being a major determinant of children's incomes in some countries, and much less important in others. To date, much of the existing research on intergenerational mobility has focused on developed nations. While this has the advantage that data sources are generally more reliable, developed nations also tend to be politically stable and to have experienced modest rates of growth.

By contrast, the growth experience of China over the past three decades has been nothing short of unprecedented. As the world's most populous nation, Chinese living standards have risen twelvefold since 1979. These rapid economic changes have also been accompanied by dramatic social transformations. All this makes China a unique case study through which to better understand the relationship between societal change and income mobility.

In traditional Chinese society (prior to 1949), most social welfare was familial. There was a strong reciprocal relationship between parents and children. Parents

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normally invested a large proportion of their income and assets in their offspring's education and career development (and the parental social network played an important role in children's access to education and the labor market). Children typically lived with their parents until marriage (and in many cases, after marriage as well). In return, parents expected their children to support them in old age.

During the Maoist era (1949–77), social welfare was universally provided in urban areas, with the aim of making Chinese society more egalitarian.[1] By and large, this successfully compressed the distribution of income and wealth over the Maoist period (Meng, 2004, 2007; Benjamin *et al.*, 2005). In addition, the early socialist revolution attempted to weaken the social ties between parental and children's occupations by making it impossible for children to inherit any meaningful wealth from their parents, and opening up better opportunities for education and occupational attainment for the children of poor families (Cheng and Dai, 1995). These might have facilitated an increase in intergenerational income mobility.

However, close family ties still exist and continue to deliver economic and social advantages from one generation to another. Indeed, some government policies have even decreased mobility. For instance, the policies of intergenerational job replacement (Dingti) and internal recruitment (Neizhao), which were introduced in 1977 and then abolished in 1986, might have reduced the level of mobility in urban China (Yu and Liu, 2004). In addition, the unique household registration system initiated in Mao's era restricted geographic labor mobility, not only from rural to urban but also from small urban cities to large urban cities, and such restrictions no doubt reduced intergenerational mobility (Wu and Treiman, 2003).

The transition from a planned to a market oriented economy initiated in the late 1970s and early 1980s has moved the urban society away from the social provision of welfare back to one that relies heavily upon individual and family responsibilities (Cai *et al.*, 2006). As a consequence, income and wealth inequality in urban China have increased sharply (see, e.g., Meng, 2004, 2007). Family networks play an important role in job attainment, which, in turn, may have reduced intergenerational income mobility. Conversely, a steady increase in geographic mobility may have boosted intergenerational mobility across China as a whole (Wu and Treiman, 2003; Yu and Liu, 2004; Takenoshita, 2007).

Studies of intergenerational mobility in China appear mainly in the sociological literature. These generally focus on social stratification and political affiliation. Such studies mainly concentrate on occupational mobility and find strong intergenerational transmission in occupations and industries (Lin and Bian, 1991; Cheng and Dai, 1995; Takenoshita, 2007). In general, Communist Party members and state employees (especially government officials) have many social advantages

[1]After the Communist Party took power in 1949, the Chinese economy was divided into two parts—urban and rural economies. One important division between the two groups was the provision of social welfare. The cradle-to-grave social welfare system was only available to the urban population, while the rural population were not covered by any welfare. This division has remained, and even today it is extremely difficult for people who were born in rural areas to gain urban household registration and to access the urban social welfare system.

in obtaining entrance into university, or locating better job opportunities for their children (Lin and Bian, 1991; Walder *et al.*, 2000; Bian, 2002; Meng, 2007).

The only other study on intergenerational income mobility in urban China is Guo and Min (2008), who use the Chinese Urban Household Education and Employment Survey 2004 (UHEES 2004). They estimate the overall intergenerational income elasticity in urban China to be 0.32 (for fathers and sons) and find that education has played an important role in promoting intergenerational income mobility. Their study uses only one income observation for parents, and one income observation for children. However, the unprecedented income growth and significant structural changes noted above may have had a differential impact on Chinese birth cohorts in recent years. Consequently, parental lifetime income may differ considerably across cohorts. Using parental income at a single point in time may substantially mis-state permanent lifetime income. We, therefore, aim to advance on Guo and Min (2008) by accounting for life cycle variation of income, and exploring the possible role of parental social status in determining children's outcomes.

This paper combines UHEES 2004 with repeat cross-sectional datasets covering nearly 20 years. To preview our results, we find much higher intergenerational income elasticity than the previous estimates by Guo and Min (2008). After accounting for fluctuations in parents' incomes and life cycle variation in children's incomes, we estimate that the intergenerational elasticity in urban China is around 0.63 for father–son, 0.97 for father–daughter, 0.36 for mother–son, and 0.64 for mother–daughter. This is an extremely high level of intergenerational persistence, and implies that intergenerational mobility is much lower in China than in most developed nations.

Exploring the channels through which income is transmitted across generations, we find that in urban China parents' and children's education, occupation, industry, and Communist Party membership are all highly correlated. We find that the main channel through which parental income affects children's income is education. For example, when we control for children's education alone, the father–son income elasticity falls from 0.63 to 0.27. Other children's characteristics do not have such a marked impact on the elasticity.

The paper is structured as follows. Section 2 reports the methodology used to estimate intergenerational income elasticity. Section 3 describes the data and summarizes the statistics. Section 4 presents estimates of intergenerational income and earnings elasticities. Section 5 explores the main channels through which income is transmitted from parents to children. The final section concludes.

2. Methodology

Studies of intergenerational mobility generally estimate the association between the socioeconomic status of parents and their offspring. Becker and Tomes (1979) first suggested a log-linear intergenerational income regression to estimate the intergenerational elasticity:

$$y_{ci} = \alpha + \beta y_{pi} + \varepsilon_i \tag{1}$$

where y_{ci} is the log of children's income, and y_{pi} is the log of parents' income. The coefficient β is the intergenerational income elasticity. The larger the elasticity, the less mobility in a given society.

Ideally, intergenerational mobility calculations should estimate the elasticity of lifetime income between children and parents. However, using the observed one-year income/earnings for parents and children is found to significantly underestimate the true intergenerational elasticity of lifetime income or earnings (Solon, 1992; Mazumder, 2005; Bohlmark and Lindquist, 2006; Haider and Solon, 2006; Dunn, 2007).[2] Following the literature (see, e.g., Solon, 1992; Altonji and Dunn, 2000; Mazumder, 2005; Leigh, 2007, among many others), the empirical model estimated in this study controls for a quadratic in children's age and a quadratic in parental age. We also include indicator variables for the 16 regions that comprise our sample:

$$y_{ci} = \alpha + \beta y_{pi} + \varphi_1 A_{ci} + \varphi_2 A_{ci}^2 + \gamma_1 A_{pi} + \gamma_2 A_{pi}^2 + \delta R_{ci} + \varepsilon_i \tag{2}$$

where A_{ci} is the child's age minus 40 (and A_{ci}^2 is its square), and A_{pi} is the parent's age (and A_{pi}^2 is its square). R_{ci} is a vector of regional dummy variables capturing regional variation in prices. There is an argument that regional income variations should not be controlled for because this may be one of the important channels through which family background can affect children's outcomes. We therefore estimate equation (2) both with and without the regional controls in order to test the sensitivity of their inclusion.

To deal with the measurement error that might arise from using one year earnings/income data to proxy for lifetime earnings/income, some previous studies also take an average of income over a number of different years for parents in order to obtain a better estimate of permanent income (Solon, 1992; Lee and Solon, 2006; Nicoletti and Ermisch, 2007). However, Mazumder (2005) argues that the transitory component of income is highly persistent and even a five year average may still provide a rather poor measurement of permanent income.

Measurement error in the dependent variable (in this case log children's income) does not in itself lead to attenuation bias. However, Grawe (2006) and Haider and Solon (2006) point out that attenuation bias can be caused by using log current earnings when children are in their 20s (a point at which current income may be quite different from average lifetime income). This can lead to a downward bias in the estimated intergenerational elasticity. Similarly, late-career earnings tend to diverge from lifetime average earnings. As a result, using log current income for parents will further downward bias the estimated intergenerational elasticity (Grawe, 2006; Haider and Solon, 2006).

To address the first bias, prior researchers use log current income measured when children are aged in their 30s or 40s (a point at which current income is close to average lifetime income). To address the second bias, studies have used average

[2]For instance, Dunn (2007) finds that in Brazil, the intergenerational earnings elasticity grows tremendously with son's age, reaching a maximum for sons aged 49, before falling slightly. Father's age is found to have much less effect on mobility estimates than the son's age, and the use of fathers aged over 40 produces quite stable elasticity estimates.

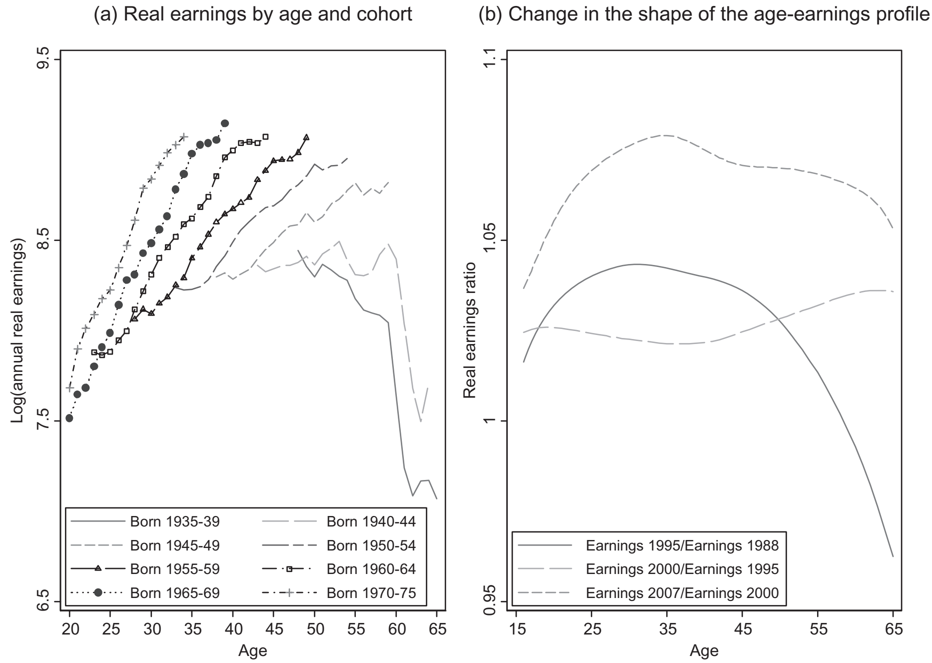

Figure 1. Change in the Level and Shape of the Age–Earnings Profile, 1986–2007
Source: Authors' own estimation results.

parental income over multiple years, or have predicted parental lifetime income by instrumenting income using parental characteristics.

In addition to general issues that arise when estimating intergenerational elasticities, the specific economic environment in China over recent decades may also pose an additional difficulty in estimating parental lifetime income. Due to rapid economic growth and structural changes that have moved China from an administratively-determined wage system to a market-oriented system, the age–earnings profile and the returns to particular demographic characteristics have shifted markedly. Figure 1 presents the changes in age–earnings profiles over the period 1988–2007. Figure 1(a) depicts the log of average real annual earnings for each particular birth cohort as they age. The different lines are for different birth cohorts, starting with those born in 1970–75 (aged 20–25 in 1995) and ending with people born in 1935–39 (aged 56–60 in 1995).

This figure may be read in two ways. In comparing different birth cohorts at a given age, it shows earnings growth across cohorts while holding age constant. We find that at the same age, the real earnings for different cohorts changed dramatically. For example, at age 40 those who were born in 1960–64 earned about 50 percent more than those born in 1950–54 earned when they were the same age. Alternatively, we can also follow the trajectory of a single line and this shows the age–earnings profile for different birth cohorts. If the age–earnings profile in China had remained constant, one would expect to see the lines remaining parallel to one another. Instead, it is evident that earnings growth has affected the wage profile differentially for different cohorts. For example, for those born in 1950–54 the profile did not flatten out until age 50, whereas for those born in 1960–64, the profile started to flatten out when they turned 38.

Another way to illustrate the same point is to look at age–earnings profiles over time. Figure 1(b) takes the average real earnings ratios for each age between 1988 and 1995, 1995 and 2000, and 2000 and 2007. When there is no change in the shape of the age–earnings profile, the graph should show horizontal lines. Figure 1(b) indicates that between 1988 and 1995, real earnings rose for workers in their 30s, but fell for workers in their 60s. This was partly reversed between 1995 and 2000 (when workers in their 50s and 60s enjoyed the most rapid earnings growth). In the period 2000–07, prime-aged workers again experienced more rapid growth in real earnings than their older and younger counterparts.

Meng *et al*. (2009) also show in their paper that the changes in the importance of age and provincial dummy variables in explaining earnings variation overtime are striking. At the beginning of the data period (the late 1980s to the early 1990s), around one-quarter of the variation in earnings could be explained by age and its squared term alone, while at the end of the period (2007), a quadratic in age explained just 2 percent. Similarly, the impact of provincial dummy variables doubled from about 10 percent to 20 percent between the late 1980s and the mid 1990s, before declining to its original level in the 2000s. Education, on the other hand, explained less than 3 percent of the variation in earnings at the beginning of the period, but over 10 percent at the end of the period.

Given the significant changes in the shape of the age–earnings profiles as well as the changing returns to demographics, it is even more problematic to assume that a single year of income can be used to proxy parents' permanent income in urban China. In this paper we address these problems using the following strategies.

First, we use OLS to estimate equation (2) with a sample of child–parent pairs where both parent and child are working in 2004. We restrict the age range for children and parents to an interval where current earnings are the best proxy for long-run earnings (we also experiment with varying this age range).

Second, aiming to address the potential downward bias caused by mis-measurement of parental income, we instrument reported parental income with parental education, occupation, and employment industry. This method, however, may suffer from certain weaknesses. If the rate of return to human capital changes significantly over time, as has occurred in China (Meng and Kidd, 1997; Zhang *et al*., 2005; Meng *et al*., 2009), we will mis-estimate the different cohorts' lifetime income. Another problem is that parental education may have an independent effect on children's income (Solon, 1992). For example, nepotism in the Chinese university admissions process might lead us to overestimate the true intergenerational income elasticity if we instrument using parental education. Similarly, nepotism in the Chinese job market might lead us to overestimate the true intergenerational income elasticity if we instrument using parental occupation.

Thus, to avoid these drawbacks associated with using single year parental income data instrumented by their education, occupation, and industry, we adopt a third method of Two-Sample-Two-Stage-Least-Squares (TS2SLS) following Arellano and Meghir (1992), Angrist and Krueger (1992), Bjorklund and Jantti

(1997), Inoue and Solon (2005), Mocetti (2007), and Nicoletti and Ermisch (2007). To implement the TS2SLS, we use a cross-sectional survey (UHEES 2004) containing information on both children's and parents' current income, education, and demographics. We then combine this with repeated cross-sectional data covering a two-decade period, and include information about parental pseudo-cohorts' earnings, education, occupation, and actual labor market experience. We predict parental permanent income using the latter data source, hoping that it will allow us to better proxy parental permanent income over a period of significant income changes.

More precisely, we use the repeated cross-section sample to estimate the following equation separately for each gender and year over the period 1986–2004:

$$y_{it} = \alpha + \beta X_{it} + \eta Prov_{it} + v_{it} \tag{3}$$

where y_{it} is the log of observed real income (deflated by province-specific price indices) for individuals who are of working age (men aged 16–60 and women aged 16–55) and working in year t. In equation (3), X_{it} is a vector of individual characteristics including years of schooling (occupation, and industry of employment), actual age and its square term in year t, while $Prov_{it}$ is a vector of provincial dummies. Since the second stage equation also includes parental age and a vector of provincial dummies, the excluded instrument is parental education (occupation, and industry of employment).

Based on the estimated coefficients from equation (3) and parental demographics from the survey that contain information on parents and children, we then predict parental income year by year for the period 1987–2004. Note that the age of the parents for the year $t - n$ for the prediction is calculated by ($A_{t-n} = A_t - n$). Because we have information from the parent–child dataset (UHEES 2004) on the exact year in which parents started work and retired, we are able to predict parental income in each of the 19 years for those who were working at that time. For parents who started the first job or retired during the 1986–2004 period, their predicted income is set to missing for the period before they started working and after they retired. After predicting parental yearly earnings, we then calculate each parent's average earnings, taking away the time trend. Only the sample of individuals who have at least five consecutive years of predicted parental income are included.

In addition to the benefit of better capturing parental permanent income, this method also significantly increases the number of observations with both children and parental incomes. However, this method may nonetheless fail to satisfy the exclusion restriction. This should be borne in mind when interpreting our results.

3. Data and Summary Statistics

The data used for this study are from two sources: the Urban Household Education and Employment Survey 2004 (UHEES) and the Urban Household Income and Expenditure Survey 1987–2004 (UHIES). The first survey was conducted jointly by the National Bureau of Statistics (NBS) and Beijing University, while the second was conducted by NBS. The UHIES is a nationwide survey (31 provinces), but due to confidentiality restrictions, we have access to data for only 16 provinces. There are 12 provinces included in the UHEES survey (Beijing, Shanxi, Liaoning, Heilongjiang, Zhejiang, Anhui, Hubei, Guangdong, Sichuan, Guizhou, Shaanxi, and Gansu). When using UHIES to estimate earnings equations as a base for predicted parental earnings, we use data for all 16 provinces we have access to (not just the 12 in the UHEES).

The 2004 UHEES collected detailed information on demographic characteristics, educational attainment, labor market status, labor market history, party membership, annual income, and annual earnings in 2004 for all household members residing in the household as well as *non-residing parents* of the household head and his or her spouse. Non-residing *children* of the household head and spouse are not surveyed. The fact that the UHEES also surveys non-resident parents makes it particularly well-suited to our empirical analysis. For parents who are retired or deceased, the survey records their last occupation and industry. The 2004 UHEES covers 9,994 urban households and 67,132 individuals, and the information on non-residing parents is reported by their children.[3]

The UHIES is a repeated cross-section dataset for the years 1987–2004, which we use to predict parental permanent income. It includes information on household income, as well as individuals' age, gender, education, occupation, and industry. We use 16 out of 31 province datasets available to us. For a detailed description of the UHIES data, see Meng *et al*. (2009).

The UHEES data can be reorganized into child–parent pairs, where each pair includes individual information for children and parents. There are two different kinds of children in the sample. Some are children residing in their parents' home where a parent is the household head (we call these the "parent-headed sample"). Another group are children residing in their own home (we call these the "child-headed sample").

Unsurprisingly, these samples have a different age distribution, with both children and parents being much younger in the parent-headed sample than in the child-headed sample. The child-headed sample on average has a higher income level for children and slightly lower income level for parents than the parent-headed sample. The mean income for children in parent-headed and child-headed samples is 13,195 and 15,457 yuan, respectively, with the t-ratio for the difference being 0.80 (and therefore not statistically significant at conventional levels). The mean income for parents in the two samples is 17,706 and 17,313 yuan, respectively and the t-ratio for the difference is 0.13 (also statistically insignificant at conventional levels). We hope that combining the two samples will give us a sample of parents and children who are reasonably representative of the general population.

The UHEES data includes 28,729 child–parent pairs. Excluding pairs with children younger than 16 or currently at school, and those with an intergenerational age difference below 14 years, there are 18,596 child–parent pairs. We further restrict our sample to those who are working, have a positive income in 2004, and with a father no older than 74 in 2004, and a mother no older than 69 in

[3]Note that because information on non-residing parents is reported by their children, these data may be noisier than the data on residing parents. To the extent that the intergenerational income mobility may differ between the two types of households, our estimates may be biased and our interpretation of the results should be treated with caution.

2004.[4] With these further restrictions and excluding missing values and very few outliers, 5475 child–father and 3431 child–mother pairs remain in the sample. For some of our specifications, we restrict the sample to those where both parent and child are within the working-age range and were working in 2004, and with non-missing income data. This further reduces the sample size to 1813.[5]

The summary statistics of children's and parents' ages, years of education, income, and earnings are reported in Table 1.[6] The top panel reports all child–parent pairs in our UHEES sample, while the lower panel reports the sample in which both children and parents are working in 2004. Below we only discuss in detail the bottom panel results, where children are between 17 and 41 years old, fathers between 39 and 60, and mothers between 37 and 55. The average age is approximately 26 for children, 54 for fathers, and 50 for mothers. Slightly more than half of the sample children are males. The average number of years of schooling is about 13.5 for children and 10 for parents. The table also shows that (conditional on working) fathers have the highest income and earnings on average, followed by children and then mothers.

Relative to the full sample, the sample with both children and parents working (lower panel of Table 1) has on average younger children (27 vs. 34) and parents (54 vs. 62). Nevertheless, the gender composition and average years of schooling do not differ across the two samples. With regard to actual income, children in the total sample on average have a higher income than children from the restricted sample. We do not compare current income for parents across the two samples as many parents from the total sample were no longer working at the time of the survey.

4. Estimated Intergenerational Income Elasticity

The intergenerational income elasticity is estimated using equation (2) for father–son, father–daughter, mother–son, and mother–daughter, and for the different age groups of children separately. We use the three different methods discussed in Section 2. The results from the first method (OLS using just the 2004 UHEES) are presented in panel A of Table 2. Controlling for child and parent age and their square terms, we estimate a father–son income elasticity of 0.24. Restricting children's age to above 23 and further to above 30, the elasticities increase to

[4]This is because when using UHIES data to predict for parental earnings, we restrict for each year that father's and mother's age is no greater than 60 and 55, respectively, and they should have at least 5 years of predicted earnings to be included in the sample. The earliest data we have for UHIES is 1987, and in order to have at least 5 years of predicted earnings the parents have to be no older than 60 or 55 in 1990 (no older than 74 and 69 in 2004).

[5]It is noticeable that while UHEES data have the greater advantage of recording both co-residing and non-residing parents and the sample size is generally quite large (nearly 10,000 households), once all the restrictions on data are in place, the remaining sample size of child–parent pairs who were both working at the time of the survey becomes quite small (1,813 pairs). This raises an issue of representativeness of our results and we acknowledge the potential drawback when interpreting the results using this small sample. However, when applying the TS2SLS method to predict for parental permanent income, many parents who were excluded as a result of not working at the time of the survey are re-included in the sample; thus our final sample of child–parent pairs using TS2SLS is quite large (8,906 pairs) and the estimated results from this sample should represent the general situation.

[6]The incomes in this study are in 2004 prices, based on provincial urban CPI indices.

TABLE 1
Summary Statistics, UHEE 2004

Total Sample		Father-Child Mean	Father-Child Std. Dev.	Mother-Child Mean	Mother-Child Std. Dev.
No. of Pairs		5,475		3,431	
Children	Age	34.16	7.14	31.75	6.76
	Age range	17–56		17–53	
	Sex (Male = 1)	0.51		0.50	
	Years of schooling	13.15	2.31	13.46	2.18
	Annual income (Yuan)	15,455	12,782	15,613	12,930
Parents	Age	62.3	7.02	57.34	6.28
	Age range	38–73		37–68	
	Years of schooling	10.06	2.31	9.88	3.64
	Predicted average annual income of 1986–2004 (Yuan)	7,628	2,295	6,000	2,199
Restricted sample (both children and parents are working)		**Father-Child Mean**	**Father-Child Std. Dev.**	**Mother-Child Mean**	**Mother-Child Std. Dev.**
No. of pairs		1,238		507	
Children	Age	26.51	3.91	24.94	3.25
	Age range	17–41		17–39	
	Sex (Male = 1)	0.52		0.54	
	Years of schooling	13.4	2.56	13.37	2.53
	Annual income (Yuan)	14,067	13,155	13,441	13,643
	Annual earnings (Yuan)	13,634	12,593	13,173	13,235
Parents	Age	53.76	3.79	49.98	3.38
	Age range	39–60		37–55	
	Years of schooling	10.24	2.96	9.98	2.73
	Annual income (Yuan)	17,642	13,513	13,849	10,801
	Annual earnings (Yuan)	16,903	12,235	13,372	10,687

0.25 and 0.32, respectively. However, when we restrict the sample to children aged 30 and above, we have relatively few observations with which to estimate mother–son and mother–daughter elasticities, and the resulting income elasticities are imprecisely estimated. Therefore, these results are not reported in the table.

If we exclude provincial dummy variables from the regression, the estimated elasticities are slightly higher in all cases. A similar pattern is found for father–daughter, mother–son, and mother–daughter elasticities. In general, the parent–son elasticities are higher than the parent–daughter elasticities.[7] For the sample of all children, the R-squared statistics range between 0.25 and 0.35, suggesting that parental income and the other demographic controls in our regression can explain up to one-third of the variation in children's incomes. Note that we do not report

[7]Some caution, though, is needed in interpreting the results for mother–son, mother–daughter, and father–daughter pairs due to a lower labor force participation rate for females, and hence higher level of sample selection for them when the estimation only includes individuals with positive incomes. If there is a systematic difference in family background between women who participate in the labor market with their counterparts who do not, estimation which only includes participants might bias the estimated intergenerational income elasticity estimated when women (mother or daughter) are involved.

TABLE 2
Intergenerational Income Elasticities Using Current Incomes

	Sons		Daughters	
Panel A: OLS	Father-Son	Mother-Son	Father-Daughter	Mother-Daughter
Log(parent's income)	0.241	0.302	0.215	0.174
	[0.0330]***	[0.0840]***	[0.0458]***	[0.0689]**
Observations	646	313	592	262
R-squared	0.301	0.348	0.254	0.303
Children's age range	17–39	17–39	17–41	17–43
Father's age range	39–60	37–55	42–60	40–55
	Sons		Daughters	
Panel B: IV = education	Father-Son	Mother-Son	Father-Daughter	Mother-Daughter
Log(parent's income)	0.346	0.451	0.301	0.355
	[0.123]***	[0.131]***	[0.116]***	[0.144]**
Observations	646	254	592	262
Children's age range	17–39	23–43	17–42	17–43
Father's age range	39–60	41–55	42–60	40–55
Panel C: IV = education & occupation				
Log(parent's income)	0.297	0.553	0.286	0.240
	[0.123]***	[0.122]***	[0.099]***	[0.107]**
Observations	651	317	594	262
Children's age range	17–39	17–43	17–42	17–43
Mother's age range	39–60	37–55	42–60	40–55
Panel D: IV = education & occupation & industry				
Log(parent's income)	0.251	0.424	0.229	0.189
	(0.072)***	(0.112)***	(0.090)**	(0.100)*
Observations	651	317	594	262
Children's age range	17–39	17–43	17–42	17–43
Mother's age range	39–60	37–55	42–60	40–55

Notes: Robust standard errors in brackets. *significant at 10%; **significant at 5%; ***significant at 1%.

Father's income and mother's income are single-year incomes (measured in 2004). Each elasticity is from a separate regression.

R-squared statistics are reported for OLS, but not for IV (since the latter is not a particularly meaningful statistic).

the R-squared in subsequent specifications, since it is not a particularly meaningful statistic in instrumental variable regressions.

Panels B–D of Table 2 present the results using one cross-sectional survey (the 2004 UHEES), but instrumenting parental income using parental demographics. Specifically we use three different combinations of instruments: education only; education and occupation; and education, occupation, and industry. The estimated elasticities increase somewhat, but the general pattern does not change much. Again, given that the sample size in the mother–son and mother–daughter samples is extremely small, plus the possible sample selection problem mentioned above, we do not place much weight on these estimated intergenerational elasticities.

We then move on to use predicted parental income from the UHIES data to estimate equation (2). Before we do so, we report some basic statistics of the predictions of the parental permanent income and examine briefly the relationship between predicted parental permanent income and their observed income in 2004 for a sample of parents who are working in 2004 and have reported a positive income.

Based on data availability, on average fathers' permanent incomes are predicted using 13 years of data, while mothers' permanent incomes are predicted using 12 years of income data. Only 5 percent of the sample has 5 or fewer years of income data, while many have 10 years or more. These data should give us a fairly good measure of the permanent income. As shown by Mazumder (2005), including more years of data reduces the downward bias due to persistent transitory shocks and corrects for the age-related errors-in-variables bias. In addition, as discussed in Section 2, there have been significant changes in the earnings levels and earnings determination mechanism. Using average predicted earnings from repeated cross-sectional data for the past 18 years helps account for the non-linear impact of these changes in earnings, thereby helping to reduce errors-in-variables bias.

We then examine the relationship between predicted parental permanent incomes and their reported income in 2004 for a group of parents who were working in 2004 and had positive income. There is a strong positive relationship between log of current and predicted permanent income (the correlation coefficient is 0.46 for fathers and 0.55 for mothers). For both fathers and mothers, the standard deviation of the difference between the current and permanent incomes is about 0.6, suggesting that for two-thirds of respondents, the difference between their permanent and current incomes is less than 60 log points (about 80 percent). This is a larger gap than one would expect to observe in a developed nation, and can be explained by the rapid changes in the Chinese labor market over recent decades, which (as noted above) have had quite differential impacts across the working-age population.

Having examined the reliability of the predicted permanent income data for parents, we use these data to estimate equation (2) and report the results in Table 3. Columns 1–4 in panels A and B of Table 3 show father–child and mother–child intergenerational income elasticities for different age groups of children, while column 5 restricts the sample to individuals whose fathers were working in 2004. Thus, the permanent income elasticities reported in column 5 are essentially comparable to the current income elasticities reported in Table 2 (which use current year reported incomes for both children and parents). In particular, they are comparable to the results reported in panel D of Table 2 where parental education, occupation, and industry are used as the instruments.[8]

The father–son income elasticities estimated using predicted permanent fathers' incomes are double those estimated using current year reported income. For example, for the total sample the OLS estimated intergenerational income elasticity using 2004 current income is 0.24 (Table 2, panel A, column 1).

[8]The sample sizes in Table 2 are slightly smaller than those shown in column 5 of Table 3, panel A, since there are a handful of parents who are working but for whom we do not observe incomes. Dropping these cases makes no substantive difference to the comparison—the elasticity is much higher when we use permanent parental incomes (Table 3) than when we use one-year parental incomes (Table 2).

TABLE 3

Intergenerational Income Elasticities Using Predicted Parental Permanent Incomes

Panel A:	All Children (1)	Children Aged >= 23 (2)	Children Aged >= 30 (3)	Children Aged 30–42 (4)	Child and Parents Both Working in 2004 (5)
Father-son					
Log(father's income) [IV = edu.&occ. & ind.]	0.563	0.549	0.654	0.634	0.468
	[0.086]***	[0.088]***	[0.110]***	[0.117]***	[0.162]***
Observations	2813	2721	2011	1638	650
Children's age range	17–58	23–58	30–58	30–42	17–39
Father's age range	38–74	40–74	48–74	48–74	39–60
Father-daughter					
Log(father's income) [IV = edu.&occ. & ind.]	0.786	0.765	0.967	0.973	0.419
	[0.086]***	[0.087]***	[0.114]***	[0.121]***	[0.159]***
Observations	2662	2573	1896	1593	595
Children's age range	17–54	23–54	30–54	30–42	17–41
Father's age range	40–74	40–74	47–74	47–74	42–60
Panel B:	All Children	Children Aged >= 23	Children Aged >= 30	Children Aged 30–42	Child and Parents Both Working in 2004
Mother-son					
Log(mother's income) [IV = edu.&occ. & ind.]	0.356	0.350	0.366	0.357	0.567
	[0.071]***	[0.072]***	[0.090]***	[0.094]***	[0.140]***
Observations	1734	1648	1014	909	352
Children's age range	17–53	23–53	30–53	30–42	17–39
Mother's age range	37–69	41–69	49–69	49–69	37–55
Mother-daughter					
Log(mother's income) [IV = edu.&occ. & ind.]	0.504	0.500	0.632	0.636	0.370
	[0.077]***	[0.078]***	[0.103]***	[0.106]***	[0.169]**
Observations	1697	1616	1015	944	307
Children's age range	17–50	23–50	30–50	30–42	17–38
Mother's age range	40–69	42–69	48–69	48–69	40–55

Notes: Robust standard errors in brackets. *significant at 10%; **significant at 5%; ***significant at 1%.

Instrumenting fathers' income with education, occupation, and industry, the elasticity is 0.25 (Table 2, panel D, column 1). However, if we use fathers' predicted *permanent* income with the same instruments, the intergenerational elasticity for the same sample (in which both parents and children worked in 2004) rises to 0.47 (Table 3, Panel A, column 5). For the total sample, the estimated elasticity is even higher at 0.56 (Table 3, Panel A, column 1).[9]

For father–daughter pairs, the increase in elasticity when using fathers' permanent income is quite similar. The estimated elasticity increases from the OLS estimate of 0.22 (in column 3, panel A of Table 2), to the IV estimate of 0.23 (column 3, panel D of Table 2), and to the TS2SLS estimate of 0.42 (column 5 in the bottom part of panel A of Table 3). If we consider the total sample, the estimated father–daughter elasticity using predicted fathers' permanent income is even higher at 0.79. For mother–son and mother–daughter pairs, the increase in estimated elasticities from using mothers' permanent income is not as large as for father–son and father–daughter pairs, but the general pattern is consistent. These results, however, suffer more from sample selection problems due to the lower labor force participation of women (and hence are less reliable).

Our preferred specification restricts the sample to children aged 30–42 (to account for lifecycle bias), and uses parental predicted permanent income. The estimated elasticities for the total sample in this specification are 0.63 for father–son, 0.97 for father–daughter, 0.36 for mother–son, and 0.64 for mother–daughter.

The above results use predicted parental permanent income with education, occupation, and industry as the instruments, and including provincial dummies. We also estimated the same regressions without provincial controls. These results are reported in Appendix A. We find that excluding provincial dummy variables results in slightly higher intergenerational income elasticities across the board.

We also estimate the same regressions using predicted parental permanent income with different combinations of education, occupation, and industry variables as the instruments. The results are quite similar to those obtained in Table 3. In general, using education alone as the instrument results in slightly higher intergenerational elasticity estimates. In attempting to see what explains the difference, we found that the unconditional correlation between child income and parental income is very similar using either instrument set (see Appendix B for the comparison of the unconditional relationships). Only when we introduce regional dummy variables into the regression do the estimates diverge. Our conjecture is that perhaps within a particular region there is more income variation across education levels than across occupation/industry groups. Put another way, within a particular region, education is a better predictor of income than are occupation and industry.

[9]The results in column 1 of Table 3 and those reported in column 5 are from two different samples. The results reported in column 1 are for the total sample, while those reported in column 5 are for the sample of people where both parent and child were working at the time of the survey in 2004. Obviously, the sample with both parent and child working in 2004 is a special group because the children in that subsample were old enough to be working (instead of being at school) and the parents were young enough not to be retired. The intergenerational income elasticity for the subsample group is lower than that found for the total sample. This may be related to the fact that the children in this subsample were relatively new to the labor market and their earnings were a poor proxy for their lifetime permanent earnings.

Finally, we check to see whether restricting the sample to children whose parents have at least ten years (instead of five years) predicted permanent income affects our results. We find that when we change the sample in this way, the estimated intergenerational income elasticities increase slightly. For example, the estimated father–son elasticity for the total sample is 0.56, while for the restricted sample it increases to 0.61.[10]

Our findings are quite consistent with the literature, which shows that the longer the period used to generate parental permanent income, the lower the attenuation bias and the larger the estimated intergenerational income elasticity. For example, Mazumder (2005) finds that using two year average data for the U.S., the estimated intergenerational elasticity is 0.25 for father–son pairs. It rises to 0.61 when using 16 years' of fathers' earnings, an increase of 144 percent. Mazumder (2005) attributes the higher estimate to two factors: first, it reduces the downward bias that stems from transitory shocks; and second, it corrects for age-related errors-in-variables bias.

Cross-country comparison of intergenerational income mobility is hampered by the fact that different studies use a variety of empirical methods, and observe children at different ages. The father–son intergenerational income elasticity is the measure most commonly reported in the literature. Table 4 compares our results with some recent estimates of intergenerational elasticities (most are for fathers and sons, but some are for other family combinations).

Comparing our TS2SLS results for sons aged 30–42 with the studies that use similar methods (either IV or TS2SLS) and restrict children to those aged in their 30s and 40s, we find that our estimated father–son intergenerational income elasticity in urban China (0.63 with control for provincial variables and 0.66 without such a control) is at the upper end of the range of estimates for other countries. For example, the estimated elasticity is 0.22 for Canada (Fortin and Lefebvre, 1998), about 0.25 for Australia (Leigh, 2007), 0.25 for Japan (Lefranc *et al.*, 2008), 0.28 for Sweden (Bjorklund and Jantti, 1997), 0.41–0.49 for France (Lefranc and Trannoy, 2005; Lefranc *et al.*, 2008), 0.50 for Italy (Mocetti, 2007), and 0.40–0.60 for the United States (Solon, 1992; Mazumder, 2005). There are too few studies of father–daughter, mother–son, and mother–daughter elasticities to draw strong conclusions about how our results compare with those for other countries. However, given that these elasticities tend to be highly correlated within countries, it seems reasonable to conclude that urban China is relatively socially immobile for women as well as men.

Relative to other Asian countries, our estimate for China is also at the high end. For exampling, Lefranc *et al.* (2008) estimate intergenerational income elasticities of 0.25 for Japan, while the estimate of Ng (2007) for Singapore's parent–child income elasticity is 0.28. Again, however, country studies differ significantly depending on the sample selection rules, characteristics of the population considered, reliability of the earnings measure, and the instrument set used.

Our analysis demonstrates the importance of obtaining a measure of permanent income, and accounting for lifecycle bias, especially in an economy where the wage structure has changed significantly. If we use single-year income measures for

[10]These results are available upon request from the authors.

TABLE 4

Summary of Recent Studies on Intergenerational Income Elasticity for Different Countries

	Country	Data Year	Elasticity	Children's Age	Parents' Age	Methods	Authors
Father-son	Australia	2004	0.2–0.3	25–55	40 (assigned)	IV	Leigh (2007)
	Brazil	1996	0.80–0.83	30–39	30–50 (in 1976)	TS2SLS	Dunn (2007)
	Brazil	1996	0.52–0.58	25–64	34 years older than sons	TS2SLS	Ferreira and Veloso (2006)
	Britain	1991–2003	0.32	31–45	N.A.	IV	Nicoletti and Ermisch (2007)
	Britain	1991	0.56–0.58	33 (mean)	47.5 (mean)	IV	Dearden *et al.* (1997)
	Canada	1986–1994	0.22	30–39	when child aged 15	IV	Fortin and Lefebvre (1998)
	Canada	1998	0.21	32–35	45.5 (mean)	OLS	Corak (2001, 2006)
	China (Urban)	2004	0.32	N.A.	N.A.	OLS	Guo and Min (2008)
	Ecuador	1994	1.13	24–40	45–60	TS2SLS	Grawe (2001,04)
	France	1993	0.41	30–40	55–70	TS2SLS	Lefranc and Trannoy (2005)
	France	1985–2003	0.50	30–50	40 (assigned)	TS2SLS	Lefranc *et al.* (2008)
	Germany	1984	0.12	25 (mean)	50–51 (mean)	OLS	Couch and Dunn (1997)
	Italy	2000–2004	0.50	30–45	30–50 (in 1977–1979)	TS2SLS	Mocetti (2007)
	Japan	1985–2005	0.25	30–50	40 (assigned)	TS2SLS	Lefranc *et al.* (2008)
	Malaysia	1976/1989	0.54	>=23	24–59	IV	Grawe (2004)
	Nepal	1995	0.32	24–40	45–60	TS2SLS	Grawe (2001, 2004)
	Pakistan	1991	0.24	25–35	45–60	TS2SLS	Grawe (2001, 2004)
	Peru	1985	0.67	24–40	45–60	TS2SLS	Grawe (2001, 2004)
	South Africa	1993–1998	0.61	25.1 (mean)	53.7 (mean)	OLS	Hertz (2001)
	Sweden	1991	0.28	30–39	43.3 (mean)	TS2SLS	Bjorklund and Jantti (1997)
	Sweden	1992	0.13	25–51	52	OLS	Osterberg (2001)
	United States	1993	0.47	28–41	40.2 (mean)	IV	Grawe (2004)
	United States	1981/1984	0.39	14–59	N.A.	IV	Altonji and Dunn (2000)
	United States	1984	0.41	25–33	44 (mean)	OLS	Solon (1992)
	United States	1984	0.53	25–33	44 (mean)	IV	Solon (1992)
	United States	1987	0.52	28–36	45 (mean)	TS2SLS	Bjorklund and Jantti (1997)
	United States	1995–98 for children	0.57	27–32 (1995)	13–55 (1970)	TS2SLS	Mazumder (2005)
		1970–85 for father		30–35 (1998)	27–69 (1984)		
Father-daughter	Canada	1986–1994	0.22	30–39	when child aged 15	IV	Fortin and Lefebvre (1998)
	Britain	1991	0.64–0.66	33 (mean)	47.5 (mean)	IV	Dearden *et al.* (1997)
	United States	1995–98 for children	0.61	27–32 (1995)	13–55 (1970)	TS2SLS	Mazumder (2005)
		1970–85 for father		30–35 (1998)	27–69 (1984)		
	United States	1981/1984	0.4	14–59	N.A.	IV	Altonji and Dunn (2000)
Mother-son	United States	1981/1984	0.29	14–59	N.A.	IV	Altonji and Dunn (2000)
Mother-daughter	United States	1981/1984	0.27	14–59	N.A.	IV	Altonji and Dunn (2000)
	South Africa	1993–1998	0.66	27.7 (mean)	52.4 (mean)	OLS	Hertz (2001)
Parent-daughter	United States	1991	0.35–0.49	33.57 (mean)	39 (mean)	OLS on panel data	Chadwick and Solon (2002)
Parent-child	Singapore	2002	0.23	23–29	N.A.	OLS	Ng (2007)
			0.28	23–29	N.A.	IV	Ng (2007)
Parent-married sons	Japan	1993–2004	0.41–0.46	36.2 (mean)	60.0 (mean)	IV	Ueda (2007)
Parent married daughter	Japan	1993–2004	0.30–0.38	36.4 (mean)	57.2 (mean)	IV	Ueda (2007)
Parent single daughter	Japan	1993–2004	0.30	27.9 (mean)	55.6 (mean)	IV	Ueda (2007)

parents, we obtain substantially lower estimates of the intergenerational income elasticity for urban China (indeed, such estimates imply that urban China is an extremely socially mobile country). However, when we use predicted permanent incomes, and restrict the sample of children to those aged 30–42 (to account for lifecycle bias), we obtain intergenerational elasticities that are sometimes twice as large.

It is worth noting that (in common with some other studies) we present only intergenerational elasticities, and do not calculate intergenerational *correlations*. The relationship between the elasticity and the correlation is $\rho = \beta \frac{\sigma_p}{\sigma_c}$, where β is the elasticity, and σ_p and σ_c are the standard deviations of log income in the parents' generation and the children's generation, respectively. To calculate ρ would require information on the underlying variance in permanent incomes. However, given that income inequality in urban China (measured by annual incomes) has risen markedly over recent decades, it is likely that inequality of permanent incomes has also risen. Since permanent income is likely to be more dispersed in the 2000s than in the 1980s, it is probable that the intergenerational income correlation in urban China would be *lower* than the intergenerational income elasticity.

5. How Is Income Earning Ability Transmitted Across Generations?

In this section, we analyze how income earning ability is transferred across generations in urban China, focusing particularly on the role of education, party membership, occupation, and industry.

Education is believed to be a significant pathway for intergenerational transmission for many countries. Using our full sample, we therefore estimate intergenerational educational transmission by both schooling years and by three categories of educational attainment: (1) lower secondary schooling or less; (2) upper secondary schooling; and (3) college and above.

Appendix C cross-tabulates the education level of parents and children. The first rows of panels A and B present the educational distribution of the children. In the total sample, 28 percent of sons have lower secondary schooling or less, 34 percent have upper secondary schooling, and 38 percent have a college degree. Panel C presents parent–child correlations. Depending on which combination we look at (father–son, father–daughter, mother–son, or mother–daughter), between 44 and 49 percent of children are in the same education category as their parents. Measured in years of schooling rather than categorically, the correlation coefficient between parents' and children's education ranges from 0.22 for mothers and sons to 0.38 for mothers and daughters (see the bottom panel of Table 4).[11] Among fathers who have a college degree, 62 percent of their sons have a college degree. Among mothers with a college degree, 65 percent of their sons have a college degree. A similar pattern can be observed for daughters.

Appendix D presents the estimated relationship between parents' and children's schooling years by children's birth year and gender, with the sample

[11]These correlations are similar to those estimated for rural China by Hertz *et al*. (2007).

restricted to children aged 25 years or older (since respondents aged less than 25 are more likely to be in the process of completing their education). The chart shows that the intergenerational association of schooling is approximately three times higher for children born in the late-1970s than for children born in the early-1950s. This increase can be explained partly by the Chinese Cultural Revolution which ended in 1976, followed by the restoration of the University Entrance Examinations in 1977. For children born before 1960, the Cultural Revolution shock reduced the gap in educational attainment between those with higher-educated parents and those with less-educated parents (Meng and Gregory, 2002).

By international standards, the intergenerational education correlation in urban China is relatively low (compare our results with Hertz *et al.*, 2007). There are two main reasons for the low correlation. First, because the Cultural Revolution reduced education levels for an entire cohort, there was a disconnection between parental and children's education for cohorts where either parents or children were young adults during the Cultural Revolution. Second, average levels of educational attainment have risen considerably since the late 1990s.

Parents' social networks can play an important role in providing their children with access to better opportunities in education and the labor market (Lin and Bian, 1991; Walder *et al.*, 2000; Meng, 2007). Communist Party membership can be transferred across generations through parental role models and social networks.

The intergenerational transmission of occupations and industries is more complicated. It depends on whether the parental social network plays an important role in children's entrance to the labor market and their promotion in the workplace; whether there are entry barriers due to crafts, professional and technical skills that are handed down; whether the attitudes and norms of family ties differ between rich and poor parents; and whether cohabitation with parents strengthens intergenerational persistence through the effects on beliefs and preferences (Mocetti, 2007).

Appendix E reports the persistence matrices of Communist Party membership, occupation, and sector of employment between children and parents by gender using the total sample. The first panel of Appendix E shows that if the parents are party members, their children are 4–10 percentage points more likely to be party members (for sons, this represents about a 10–20 percent increase in the probability of joining the party, while for daughters it represents more than a 50 percent increase in the probability of party membership).

The second panel indicates a very strong persistence in occupation between children and parents. If fathers or mothers are working in the occupational category *professionals and technicians*, their children are 35–40 percentage points more likely to also be working in this occupation than those whose parents are not (this represents a near-tripling in the probability of being in this occupational category). The differences for *administration staff* are 17–23 percentage points (approximately a doubling in the probability), while for *production and transportation workers* the differences are 16–30 percentage points (having a father who was a production worker approximately triples the probability that a child will enter this occupational grouping, while having a mother who was a production worker doubles the probability).

The third panel of Appendix E presents the proportions of sons and daughters working in the state-owned sector based on whether their fathers and mothers also work in that sector. Seventy-one percent of sons and 67 percent of daughters work in the state-owned sector. Children are 6–9 percentage points (about 10 percent) more likely to work in the state-owned sector if one of their parents also worked in that sector.

Following Hertz (2008), we also conduct a decomposition to see whether income mobility is higher for children whose fathers are party members than for those whose fathers are not in the party. We find that the within-group elasticities and persistence for those whose fathers are party members are 0.40 and 0.37, respectively. For those whose fathers are not party members, the elasticities and persistence are 0.69 and 0.81, respectively. These results indicate that within-group mobility is much higher for those whose fathers are party members than those whose fathers are not (and conversely that intergenerational persistence is lower).[12]

Finally, we re-estimate intergenerational income elasticities as those reported in Table 3, but this time controlling for children's education, party membership, occupation, and industry individually and together. The results are reported in Table 5. Using the sample of children aged 30–42, we find that if we include the education level of the children into the regression, the estimated intergenerational income elasticity for father–son reduces from 0.63 to 0.27. For mother–son, father–daughter, and mother–daughter there is also similar reduction. For brevity, these results are not presented in the table but are available upon request from the authors.

Including children's occupation also reduces the intergenerational elasticities significantly but not as much as including education. Party membership and industry of employment have an even smaller impact on the intergenerational elasticities. Although these tests are somewhat crude, they suggest that the most important channel through which parental income affects children's income is education.

6. Conclusions

An old Chinese saying holds that families will "be poor no more than three generations and be rich no more than three generations." This suggests that in China, as in many other nations, there is a strong popular belief in social mobility. Our findings challenge this view. At least for modern-day urban China, we find a strikingly low level of intergenerational mobility. Our preferred estimates show that the intergenerational income elasticities are 0.63 for father–son, 0.97 for father–daughter, 0.36 for mother–son, and 0.64 for mother–daughter. Internationally, our estimated father–son elasticity places urban China among the least socially mobile places in the world.

Exploring possible pathways, we find that education, especially college study, is an important channel through which earnings ability is transmitted from parents to children (though the intergenerational association of education is still lower in

[12]After adjusting for the between-group effect the persistence changes very slightly. These results are available upon request from the authors.

TABLE 5

What Explains the Intergenerational Elasticity?

Panel 1: Father-son	*Dependent variable: Log (son's income)*							
Log (father's income) [IV = edu + occ + ind]	0.634	0.272	0.555	0.316	0.442	0.267	0.165	0.138
	[0.117]***	[0.117]**	[0.117]***	[0.117]***	[0.120]***	[0.117]**	[0.117]	[0.118]
Education of son		Yes				Yes	Yes	Yes
Party membership of son			Yes			Yes	Yes	Yes
Occupation of son				Yes			Yes	Yes
Industry sector of son					Yes			Yes
Observations	1638	1638	1637	1633	1638	1637	1632	1632
R-squared	0.24	0.30	0.26	0.30	0.28	0.30	0.33	0.35
Panel 2: Mother-son	*Dependent variable: Log (son's income)*							
Log (mother's income) [IV = edu + occ + ind]	0.357	0.120	0.310	0.159	0.244	0.115	0.049	0.056
	[0.094]***	[0.092]	[0.094]***	[0.094]*	[0.092]***	[0.092]	[0.094]	[0.093]
Education of son		Yes				Yes	Yes	Yes
Party membership of son			Yes			Yes	Yes	Yes
Occupation of son				Yes			Yes	Yes
Industry sector of son					Yes			Yes
Observations	909	909	909	906	909	909	906	906
R-squared	0.25	0.32	0.26	0.32	0.31	0.33	0.35	0.39
Panel 3: Father-daughter	*Dependent variable: Log (daughter's income)*							
Log (father's income) [IV = edu + occ + ind]	0.973	0.551	0.922	0.552	0.593	0.550	0.388	0.288
	[0.121]***	[0.121]***	[0.121]***	[0.118]***	[0.121]***	[0.120]***	[0.118]***	[0.117]**
Education of daughter		Yes				Yes	Yes	Yes
Party membership of daughter			Yes			Yes	Yes	Yes
Occupation of daughter				Yes			Yes	Yes
Industry sector of daughter					Yes			Yes
Observations	1593	1593	1590	1582	1593	1590	1579	1579
R-squared	0.20	0.29	0.22	0.29	0.29	0.29	0.32	0.35
Panel 4: Mother-daughter	*Dependent variable: Log (daughter's income)*							
Log (mother's income) [IV = edu + occ + ind]	0.636	0.349	0.609	0.355	0.451	0.352	0.240	0.231
	[0.106]***	[0.100]***	[0.105]***	[0.104]***	[0.102]***	[0.100]***	[0.101]**	[0.100]**
Education of daughter		Yes				Yes	Yes	Yes
Party membership of daughter			Yes			Yes	Yes	Yes
Occupation of daughter				Yes			Yes	Yes
Industry sector of daughter					Yes			Yes
Observations	944	944	944	939	944	944	939	939
R-squared	0.21	0.29	0.22	0.29	0.29	0.30	0.32	0.35

Notes: Robust standard errors in brackets. *significant at 10%; **significant at 5%; ***significant at 1%.
Parent income is the permanent income of the child's mother or father (see text for details). Each column is a separate regression.
Children used are aged 30–42.

urban China than in many other nations). We also estimate intergenerational correlations for parental party membership, occupation, and industry. The intergenerational occupational correlation is particularly high. However, it is also possible that factors we do not perfectly observe in our data—such as genes, health, or social networks—are also significant channels of intergenerational transmission in urban China.

The fact that we find the main channel for intergenerational income transmission is through education has important policy implications. The very low income mobility in current day urban China may reflect the need for policymakers to focus on providing more educational opportunities to bright children with less educated parents. Although the topic is beyond the scope of this paper, it points us to a future research area.

Finally, it is important to point out that our study focuses only on urban China. Given the large income gap and significant institutional differences between rural and urban China, our results may not generalize to the whole country. However, from a policy perspective, our estimated elasticity is still an important parameter. The large scale rural–urban migration which is occurring in China offers the potential for upwards social mobility for those born in rural areas. It may at the same time create the possibility for downwards social mobility for those born in urban areas.

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Supporting Information

Additional Supporting information may be found in the online version of this article:

Appendix A (without prov dummy): Intergenerational mobility elasticities using permanent parental income with education/occupation/industry as instruments

Appendix B: Unconditional Correlation between Children and Fathers' Predicted Incomes Using Education or Occupation/Industry as Instruments

Appendix C: Educational persistence and correlation across generations

Appendix D: Intergenerational Education Correlation (regression coefficients) by Children's Birth Year in Urban China, 2004

Appendix E: Intergenerational persistence in party membership, occupation and sector of employment

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