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# Estimating returns to education using twins in urban China☆

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ABSTRACT

This paper empirically estimates the return to education using twins data that the authors collected from urban China. Our ordinary least-squares estimate shows that one year of schooling increases an individual's earnings by 8.4%. If we use a within-twin fixed effects model, the return is reduced to 2.7%, but rises to 3.8% after the correction of measurement error. These results suggest that a large portion of the estimated returns to education is due to omitted ability or the family effect. We further investigate why the true return is low and the omitted ability bias high, and find evidence showing that it may be a consequence of China's education system, which is highly selective and exam oriented. More specifically, we find that high school education may mainly serve as a mechanism to select college students, but as a human capital investment per se it has low returns in terms of earnings. In contrast, both vocational school education and college education have a large return that is comparable to that found in the United States.


## 1. Introduction

Although estimating the return to education has been an important econometric exercise since the seminal work of Mincer (1974), only recently have economists begun to estimate it using Chinese data. Several studies that draw on data from urban China from the 1980s and 1990s find rather low returns, with one year of schooling increasing earnings by only 2–4% (Byron and Manaloto, 1990; Meng and Kidd, 1997). This finding has caught the attention of many labor economists, who generally think that the estimates of the return to education in China were formerly low because most of the urban economy was still under a planned regime in the 1980s and 1990s.[1] However, they believe that the return should have increased after more than two decades of economic transition from a planned regime to a market regime, as in market economies in which a large gradation in earnings according to the level of education reflects the return to the investment of individuals in education (Becker, 1993; Mincer, 1974).[2] Recent data have shown that the return to education has indeed risen in China (see e.g., Fleisher and Wang, 2005; Heckman and Li, 2004; Li, 2003). Using the urban household survey that was conducted by the National Bureau of Statistics for 1988–2001, Zhang et al. (2005) find a dramatic increase in the return to education in urban China from only 4% in 1988 to more than 10% in 2001.[3]

Despite the rapid accumulation of evidence on the return to education in China, little has yet been done to establish causality.[4] An ordinary least-squares (OLS) estimation of the effect of education on earnings cannot prove causality, because well-educated people may have high earnings as a result of greater ability or a better family background. In other words, education may be correlated with unobserved ability or family background effects, which would render any causal effect between education and earnings spurious. Due to this difficulty posed by the endogeneity arising from unobserved ability, the true return to education in China remains elusive.

Our first goal in this paper is to empirically measure the causal effect of education on earnings by using twins data that one of the authors collected in urban China. As is argued in the literature (Ashenfelter and Krueger, 1994; Behrman et al., 1996; Behrman and

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[1] See Groves et al. (1994) for a description of the planned economy and its transition and Li and Zax (2003) for a study of labor supply during economic transition.

[2] In fact, this assertion has contributed to a lively debate among social scientists, and in particular sociologists, who study institutional transformation and social stratification in former state socialist societies (see e.g. Walder (1996), Zhou (2000) and Giles et al. (2006a, 2006b)).

[3] Similar patterns have been documented in rural China (de Brauw et al., 2002; Rozelle et al., 2001; Rozelle and Swinnen, 2004) and other transition countries (Andren et al., 2005; Earle and Sabirianova, 2002; Fleisher, 2005).

[4] Heckman and Li (2004) are an exception.

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Rosenzweig, 1999; Miller et al., 1995),[5] since monozygotic (from the same egg) twins are genetically identical and have a similar family background, the effects of unobserved ability or family background should be similar for both twins. Thus, taking the within-twin difference will, to a great extent, reduce the unobservable ability or family background effects that cause bias in OLS estimations of the return to education.[6] Intuitively, by comparing the earnings of identical twins with a different number of years in education, we can be more confident that the correlation that we observe between education and earnings is not due to a correlation between education and an individual's ability or family background.

Our empirical work shows that a large part of the effect of education on earnings from the OLS estimates is in fact due to the effects of unobserved ability or family background. According to our estimate the return to one more year of education is 8.4%, which is close to other recent estimates using Chinese or Asian data (see, for example, Heckman and Li, 2004; Psacharopoulos, 1992; Zhang et al., 2005). However, once we use the within-twin fixed effects model, the return is reduced to 2.7%, which suggests that much of the estimated return that is obtained using the OLS model is due to omitted ability or family effects. In other words, education in urban China is more important in selecting youngsters of high ability to progress through the system than it is in providing knowledge or training that will enhance earnings. This finding is confirmed by a generalized least-squares estimation that includes the education of the co-twin as a covariate. Finally, we also address the potential bias that is caused by error in the measurement of the education variable by using the instrumental variable approach of Ashenfelter and Krueger (1994). The estimated return to education rises by about one percentage point to 3.8% once the measurement error has been corrected.

The low estimated return to education and high selectivity (or ability bias) differ sharply from within-twin estimates that use data from the United States (see, for example, Ashenfelter and Krueger, 1994; Behrman and Rosenzweig, 1999) and the United Kingdom (Bonjour et al., 2003), and thus our second goal is to ascertain what China is so different from these countries. Although the remaining features of a planned economy could be used to explain the low return, we provide an alternative explanation in this paper. We argue that the low return and high selectivity may be a consequence of Chinese education system. Because of the huge population aspiring for higher education and the limited number of college (and university) places, entrance to college is extremely competitive. The Chinese solution to this is examinations. Only the very talented can score high enough in the college entrance exams to advance to higher education. Thus non-tertiary education, and in particular high school education and the associated entrance exams, have become a very important selection mechanism. This explains why the ability bias is so high in our OLS estimates. Moreover, to prepare students for college entrance exams, non-tertiary education in China, and in particular high school education, is totally exam oriented, and thus adds little value in terms of general knowledge or workplace skills. Consequently, such exam-oriented high school education has a low return, which has also dragged down the overall return to education. Furthermore, the high school curriculum is fixed by the Ministry of Education and high-school students have to make an early decision as to whether to specialize in the arts or science. These institutional arrangements prevent students from obtaining a more general education that would be better rewarded in the workplace.

The twins data that we have collected allow us to test whether the Chinese education system should indeed be blamed for the low return to and high selectivity of education. To this end, we estimate the returns to different levels of education by using OLS, within-twin-pair, and IV estimations. Arguably, exam-oriented high school education should have the lowest return among all of the education levels, and the final-stage education levels, such as vocational and college education, should have higher returns because they are less exam oriented. Interestingly, these hypotheses are confirmed. We find that the return to high school education is almost zero, but the return to college education is very large. According to our estimates, the return to high school education is statistically not significantly different from zero, whereas the return is 22% for vocational high school, 23% for vocational college education, and is as high as 40% for college education. These findings suggest that going to high school does not pay unless an individual is also able to obtain a college degree.

The idea of using twins data to remove omitted ability bias excited many labor economists when it was first proposed, but its popularity waned when many twins studies found that the OLS estimates did not differ much from the within-twin estimates that controlled for omitted ability. Part of the reason for the low omitted ability bias in previous studies is that most of these studies draw on data from the US and UK, where education is not very selective. To the best of our knowledge, this is the first study of the return to education that draws on twins data from China, and is probably also the first to draw on Asian twins data. The education systems of Asian countries, and especially East Asian countries and regions such as Japan, South Korea, Taiwan, Hong Kong, and mainland China, as well as many European countries, are similar in that they all have very serious college entrance exams. Understandably, high schools in these countries or regions place a great deal of emphasis on exam-taking techniques, and thus education may be more selective in these regions than it is in the US and UK. Our study is also one of the first to use twins data from developing countries. Twins studies in developing countries are particularly interesting, because the omitted variable bias may be larger in these countries, in which liquidity constraints and family background are likely to be important determinants of both education and earnings (Brown and Park, 2002; Herrnstein and Murray, 1994; Hertz, 2003; Lam and Schoeni, 1993; Park et al., 2007).

Determining the true return to education is very important for China, which is currently undergoing a transition from a planned economy to a market economy. During the transition process, the Chinese government must reform all economic sectors, such as industry, banks, the medical system, and education. Given the limited resources that are available, the government needs to set priorities for expenditure. Our findings suggest that the true return to one year of schooling is at most 3.8%, which seems to be low relative to the high rate of return to physical investment in China.[7] However, the return is not uniform for all educational levels. The finding that vocational high school has a much larger return than high schools suggests that the government should invest more in vocational schools, which produce the technicians and floor workers that are most needed in China.

The remainder of this paper is structured as follows. Section 2 describes the estimation methods that draw on twins data. Section 3 describes the data and the variables. Section 4 empirically measures the return to education. Section 5 explains why the return to education is low in China. Section 6 concludes.

[5] The earliest attempt to look at siblings data in economics can be traced back to the dissertation of Gorseline (1932). Not content with siblings data, economists started to use twins data in the late 1970s, when the works of Behrman and Taubman (1976), Taubman (1976a, 1976b), and Behrman et al. (1977) were published. Interest in using twins data was revived with the work of Ashenfelter and Krueger (1994) and Behrman et al. (1994). See also Ma (2005) who examines the return to years of schooling but not that to schooling levels.

[6] Alternatively, one can directly control for the genetic effect and family background (see e.g., Zax and Rees (2002)), though it is generally difficult to observe these variables. Researchers have also relied on natural experiments to resolve causality problem (see e.g., Dale and Krueger (2002)).

[7] Although not directly comparable, the rate of return to investment in physical capital is as high as 20% (Bai et al., 2006).

## 2. Method

Our empirical work focuses on the estimation of the log earnings equations of twin pairs, which is given as follows.

$$y_{1i} = X_i\alpha + Z_{1i}\beta + \mu_i + \epsilon_{1i} \tag{1}$$

$$y_{2i} = X_i\alpha + Z_{2i}\beta + \mu_i + \epsilon_{2i}, \tag{2}$$

where $y_{ji}$ $(j=1,2)$ is the logarithm of the earnings of twin $j$ in family $i$ and $X_i$ is the set of observed variables that vary by family but not between twins, that is, the family background variables. $Z_{ji}$ $(j=1,2)$ is a set of variables that vary between twins. $\mu_i$ represents a set of unobservable variables that also affect earnings, that is, the effect of ability or family background. $\varepsilon_{ji}$ is the disturbance term, which is assumed to be independent of $Z_{ji}$ and $\mu_i$. The estimation of Eq. (1) with $\mu_i$ excluded results in an estimate of the effect of education that is generally biased, because $\mu_i$ is very likely correlated with $Z_i$.

A within-twin or fixed effects estimator of $\beta$ for identical twins, $\beta_{fe}$, is based on the first difference of Eqs. (1) and (2),

$$y_{1i} - y_{2i} = (Z_{1i} - Z_{2i})\beta + \epsilon_{1i} - \epsilon_{2i}. \tag{3}$$

The first difference removes both the observable and unobservable family effects, that is, $X_i$ and $\mu_i$. As $\mu_i$ has been removed, we can apply the OLS method to Eq. (3) without worrying about the bias that arises from the omitted ability and family background variables.

A second approach to overcoming the omitted variable bias is to directly estimate both the bias and the education effect using the generalized least-squares (GLS) method. This method was developed by Ashenfelter and Krueger (1994) for twins studies. In this approach, the correlation between the unobserved family effect and the observables is given as

$$\mu_i = Z_{1i}\gamma + Z_{2i}\gamma + X_i\delta + \omega_i, \tag{4}$$

where we assume that the correlations between the family effect $\mu_i$ and the characteristics of each twin $Z_{ji}$ $(j=1,2)$ are the same, and that $\omega_i$ is uncorrelated with $Z_{ji}$ $(j=1, 2)$ and $X_i$. The vector of the coefficients $\gamma$ measures the selection effect that is related to the family effect and individual characteristics, including education.

The reduced form for Eqs. (1), (2), and (4) is obtained by substituting Eq. (4) into Eqs. (1) and (2) and collecting the terms as follows.

$$y_{1i} = X_i(\alpha + \delta) + Z_{1i}\beta_2 + (Z_{1i} + Z_{2i})\gamma + \epsilon'_{1i} \tag{5}$$

$$y_{2i} = X_i(\alpha + \delta) + Z_{2i}\beta_2 + (Z_{1i} + Z_{2i})\gamma + \epsilon'_{2i}, \tag{6}$$

where $\epsilon'_{ji} = \omega_i + \epsilon_{ji}$, $(j=1,2)$. Eqs. (5) and (6) are estimated using the GLS method, which is the best of the estimators that allow cross-equation restrictions on the coefficients. Although both the fixed effects and GLS models control for ability and can produce unbiased estimates of the education effect $\beta$, the GLS model also allows the estimation of the selection effect $\gamma$.

## 3. Data

The data that we use are derived from the Chinese Twins Survey, which was carried out by the Urban Survey Unit (USU) of the National Bureau of Statistics (NBS) in June and July 2002 in five cities in China. Based on existing twins questionnaires from the United States and elsewhere, the survey covered a wide range of socioeconomic information. The questionnaire was designed by two of the authors of this paper in close consultation with experts on twins studies and Chinese experts from the NBS. Adult twins aged between 18 and 65 were identified by the local Statistical Bureaus through various channels, including colleagues, friends, relatives, newspaper advertising, neighborhood notices, neighborhood management committees, and household records from the local public security bureau. Together, these channels created a roughly equal probability of contacting all of the twins in these cities, and thus the sample that was obtained is approximately representative of twin pairs who reside in the same city. The questionnaires were completed through household face-to-face personal interviews. The survey was conducted with considerable care. One of the authors made several site checks of the survey work and closely monitored the data input process.

This is the first socioeconomic twins dataset compiled in China, and perhaps the first in Asia. The dataset includes rich information on the socioeconomic situation of respondents in the five cities of Chengdu, Chongqing, Harbin, Hefei, and Wuhan. Altogether there are 4683 observations, of which 3012 are from twins households. We can distinguish whether the twins in the sample are identical (monozygotic, MZ) or non-identical (fraternal, DZ). We consider a pair of twins to be identical if both twins responded that they have identical hair color, looks, and gender.[8] Completed questionnaires were collected from 3002 individuals, of which 2996 were twin individuals and 6 were triplet individuals. From these 3002 individuals, we have 914 complete pairs of identical twins (1828 individuals). We have complete information on earnings, education, and other variables for both twins in the pair for 488 of these pairs (976 individuals). The summary statistics of the identical twins and all of the twins together are reported in the first two columns of Table 1.

For the purposes of comparison, data on non-twin households in the five cities were taken from regular households on which the Urban Survey Unit (USU) conducts regular monthly surveys of its own. The USU started regular monthly surveys in the 1980s. Their initial samples were random and representative, and they have made every effort to maintain these good sampling characteristics. However, their samples have become less representative over time. In particular, because of an increasingly high (low) refusal rate among young (old) people, the samples have gradually become biased toward the oversampling of older people. The survey of non-twin households was conducted at the same time as the twin survey, and the same questionnaire was used. The summary statistics of our non-twins sample are reported in the third column of Table 1.

To examine the extent to which our twins sample is representative, we compare the identical twins sample to the other samples that we have. To facilitate such comparisons, we also provide the basic statistics for a large-scale survey that was conducted by the USU of the NBS as a benchmark (henceforth the NBS sample, reported in column 4 of Table 1).[9] Column 1 shows that 60% of our identical twins were male, and on average the twins were 35 years old, had 12 years of schooling, and had spouses who also had an average of 12 years of schooling. They had worked for an average of 15 years, and had monthly average earnings of 888 yuan, including wages, bonuses, and subsidies. The individuals in the identical twins sample were younger than those in the NBS sample and also earned less. Finally, the individuals in the non-twins sample (column 3) were older than those in the NBS sample and the twins samples.

The within-twin-pair estimation method that is used for this study controls for the family-level effects of any unobserved characteristics that may have led to the selection of twins pairs into the sample

[8] The proportion of MZ twins in our sample (60%) is close to the proportion (56%) of white male twins among U.S. veterans born between 1917 and 1927 in the NAS-NRC twins sample that is used by Behrman et al. (1994). In the sample of Australian twins that is used by Miller et al. (1995), the proportion of MZ twins is 51%. However, as our method of identifying MZ twins is not a scientific test, it is not perfect, and thus to the extent that our sample of MZ twins may contain some DZ twins, the magnitude of the ability bias may be understated.

[9] The NBS has been conducting an annual survey of urban households randomly drawn from 226 cities (counties) in China since 1986. It is the best large-scale survey of this kind.

**Table 1**
Descriptive statistics of the twins and non-twins samples.

| Variable | MZ twins (1) | All twins (2) | Non twins (3) | NBS sample (4) |
|---|---|---|---|---|
| Education (years of schooling) | 12.22 (2.89) | 12.16 (2.91) | 11.73 (3.07) | 11.62 (2.83) |
| High school dummy | 0.27 (0.44) | 0.25 (0.43) | 0.30 (0.46) | – |
| (=1 if a high school diploma was the final qualification, =0 otherwise) | | | | – |
| Vocational school dummy | 0.34 (0.47) | 0.35 (0.48) | 0.32 (0.47) | – |
| (=1 if a vocational school diploma, at either high-school or college (2-year) level, was the final qualification, =0 otherwise) | | | | – |
| College dummy | 0.13 (0.33) | 0.12 (0.33) | 0.10 (0.30) | – |
| (=1 if a university degree was the final qualification, =0 otherwise) | | | | – |
| Age | 34.78 (9.64) | 33.77 (9.22) | 43.27 (8.42) | 40.80 (11.98) |
| Gender (male) | 0.60 (0.49) | 0.59 (0.49) | 0.48 (0.50) | 0.55 (0.50) |
| Married | 0.66 (0.47) | 0.64 (0.48) | 0.94 (0.24) | – |
| (=1 if being married, =0 otherwise) | | | | – |
| Tenure | 15.03 (9.93) | 14.03 (9.50) | 21.70 (9.05) | 18.45 (12.94) |
| (the number of years in full-time work since the age of 16) | | | | |
| Earnings | 887.85 (517.91) | 872.52 (546.00) | 845.84 (549.08) | 1062.92 (840.09) |
| (monthly wages, bonuses, and subsidies in RMB) | | | | |
| Spousal education | 11.64 (3.11) | 11.69 (3.08) | 11.49 (3.49) | – |
| (years of education of spouse) | | | | – |
| Sample size | 976 | 1620 | 1277 | 23,288 |

Notes: The mean and standard deviation (in parentheses) are reported in the table. For the MZ twins sample, we restrict the sample to those twin pairs (488 pairs) for which we have complete information on earnings, age, gender, years of education, job tenure, and marital status for both twins in the pair. The NBS sample is based on a large-scale survey by the National Bureau of Statistics in six provinces

(Behrman et al., p. 677, 1996). The within-twin estimator eliminates the selection that is common to the twins, in our case the selectivity that is associated with twins being born in households that are different (e.g., in terms of parental education). If one uses the traditional Heckman lambda term to correct for selection through being a twin, it would be the same for both of the twins in terms of common background variables. However, selection at the individual level (e.g., being in the labor force) cannot be accounted for by the within-twin difference. In the literature on twins studies, there is no good way to address the selection at the individual level directly. We deal with this problem partially by using a larger sample with earnings of the last job for those were not working currently (Behrman et al., 1996), and examine the sensitivity of our main results to this treatment in Section 5.2.

To ensure the good performance of the within-twin estimation of the return to education, the within-twin variation in education needs to be of a sufficient size. We check the within-twin variation in education and find it to be fairly large. Fifty-three percent of the twin pairs had the same education, 13% had one year's difference in education, about 10% had two years' difference, and the remaining 24% had a difference of more than two years. Moreover, more than half of the twin pairs have different educational levels (Table 2). Taking high school completion as an example, in 82 of the twin pairs both twins have a high school degree. However, for a larger number of the twin pairs (87), one twin has a high school diploma as the final qualification and the other twin has a different final qualification (37 have junior high school education, 21 have vocational high school education, 20 have vocational college education, and 9 have college education). These numbers suggest that we have a large variation of within-twin difference in education that is sufficient for the fit of the regressions.

## 4. Return to education

In this section, we report the estimated return to education using different samples and methods. We start with the OLS regressions using the whole sample, including MZ twins and non-twins, and then conduct the same OLS estimation using the MZ twins sample only and compare the estimated coefficients to those that are estimated using the whole sample. This comparison may serve as a means of checking the representativeness of the MZ twins sample. We then conduct the within-twin fixed effects and GLS estimations using the twins sample, and finally examine the possible bias in the fixed effects estimates.

### *4.1. OLS regressions using the whole sample*

In the first two columns of Table 3, we report the results of the OLS regressions using the whole sample, including both MZ twins and non-twins. The dependent variable is the logarithm of monthly earnings, and the standard errors that are reported in parentheses are robust to heteroscedasticity and clustering at the family level. In column 1, we show a simple regression with education, age, age squared, and gender as the independent variables. This regression shows the return to education to be quite large. One more year of schooling increases an individual's earnings by 6.6%, which is quite precisely estimated with a standard error of 0.004. The positive coefficient of age and the negative coefficient of age squared are both significant at the 10-percent level. Earnings increase with age before

**Table 2**
Distribution of education levels of the MZ twins sample (488 pairs of MZ twins).

| Twin 1 | Twin 2 | | | | |
|---|---|---|---|---|---|
| | Junior high school | High school | Vocational high school | Vocational college | College |
| Junior high school | 90 | | | | |
| High school | 37 | 82 | | | |
| Vocational high school | 20 | 21 | 38 | | |
| Vocational college | 21 | 20 | 19 | 62 | |
| College | 2 | 9 | 6 | 33 | 28 |
| Total | 170 | 132 | 63 | 95 | 2 |

**Table 3**
OLS and fixed effects estimates of the return to education for twins and non-twins in urban China (dependent variable: log earnings).

| Sample model | All | | MZ twins | | | | | |
|---|---|---|---|---|---|---|---|---|
| | OLS (1) | OLS (2) | OLS (3) | OLS (4) | FE (5) | FE (6) | GLS (7) | GLS (8) |
| Own education | 0.066$^{***}$ (0.004) | 0.066$^{***}$ (0.004) | 0.083$^{***}$ (0.006) | 0.084$^{***}$ (0.006) | 0.025$^{*}$ (0.015) | 0.027$^{*}$ (0.015) | 0.025$^{**}$ (0.012) | 0.027$^{**}$ (0.012) |
| Sum of education | | | | | | | 0.033$^{***}$ (0.007) | 0.033$^{***}$ (0.007) |
| Age | 0.023$^{**}$ (0.009) | 0.011 (0.012) | 0.043$^{***}$ (0.015) | 0.038$^{**}$ (0.019) | | | 0.040$^{***}$ (0.014) | 0.034$^{**}$ (0.016) |
| Age squared | −0.021$^{*}$ (0.012) | −0.023$^{*}$ (0.014) | −0.048$^{**}$ (0.022) | −0.053$^{**}$ (0.024) | | | −0.043$^{**}$ (0.020) | −0.049$^{**}$ (0.021) |
| Gender (male) | 0.225$^{***}$ (0.024) | 0.218$^{***}$ (0.024) | 0.200$^{***}$ (0.037) | 0.198$^{***}$ (0.037) | | | 0.203$^{***}$ (0.037) | 0.201$^{***}$ (0.037) |
| Married | | −0.024 (0.042) | | −0.020 (0.050) | | −0.043 (0.052) | | −0.019 (0.043) |
| Tenure | | 0.015$^{***}$ (0.003) | | 0.010$^{*}$ (0.006) | | 0.015 (0.010) | | 0.012$^{**}$ (0.006) |
| Twin pairs | | | | | 488 | 488 | 488 | 488 |
| Observations | 2253 | 2253 | 976 | 976 | 976 | 976 | 976 | 976 |
| R-squared | 0.15 | 0.16 | 0.22 | 0.22 | 0.01 | 0.02 | | |

Notes: Standard errors in parentheses are robust to heteroscadesticity and for OLS and GLS clustering at the family level.
* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

55 years of age, but start to drop after that. The gender difference in earnings is quite large in urban China, with men earning 22.5% more than women.

When we add other control variables in the second column, including marital status and tenure, the estimated coefficient of education remains unchanged, which suggests that omitting these variables results in no bias in the estimated return to education. We do not find a marriage premium in the sample, as the marriage dummy is not significant at the conventional level. Job tenure has a positive effect, with one more year in a post increasing earnings by 1.5%.

### 4.2. OLS regressions using the monozygotic twins sample

In this subsection, we repeat the same OLS regressions using the MZ sample. Comparing the OLS results for the MZ twins sample with those for the whole sample provides a means of checking the robustness of the estimated coefficients using different samples. As we only use the MZ twins sample, the sample size is reduced to 976 observations (or 488 pairs of twins).

The regression results that are reported in the third and fourth columns of Table 3 suggest that the return to education is larger for our MZ twins sample. The return to education is 8.3% for the simple regression in column 3, and becomes even larger when other control variables are included in column 4.[10] Thus, the OLS estimate of the return to education for the twins sample is about 1.7–1.8 percentage points more than that for the whole sample. The estimated coefficients of most of the other variables are very similar for both samples.

To summarize, the OLS estimate of the return to education is rather large, and even after we control for many covariates the remaining effect is 0.084 (column 4). However, we still do not know how much of this effect is the true return to education, and how much is due to the effects of unobserved ability or family background. In the following, we thus use within-twin and GLS estimations to remove the unobservables and estimate the true return.

### 4.3. Within-twin and GLS estimations

In columns 5 and 6 of Table 3, we report the results of the within-twin fixed effects estimation, or the estimation using Eq. (3). As MZ twins are of the same age and gender, these variables are dropped when calculating the first difference.

The within-twin estimation shows that much of the return to education that is obtained by the OLS estimation is the result of the effects of unobserved ability or family background. Note that the within-twin estimate of the return to education is much smaller than the OLS estimate. Taking column 6 as an example, the education effect is 0.027, which is only about one third of the OLS estimate using the same twins sample. This suggests that two thirds of the OLS estimate of the return is actually due to the unobserved ability or family effect,[11] although it should be noted that these estimates may be biased by measurement error, which we will address presently. The other control variables are not significant in the within-twin estimation.

We next turn to the GLS estimator for Eqs. (5) and (6), which can directly estimate both the return to education and the ability or family background effect. The GLS estimates are reported in the last two columns of Table 3, including the covariates that are used in the OLS estimates. In addition to an individual twin's own education, we also include the sum of the education of both twins as an independent variable. The coefficient of this new variable is thus the estimated ability or family effect, that is, $\gamma$ in Eqs. (5) and (6). The GLS model is estimated by stacking Eqs. (5) and (6) and fitting them using the SURE model.

The GLS estimation again shows that the return to education is small, whereas the omitted ability or family effect is large. The coefficients of an individual's own education are only 0.025–0.027, which are exactly the same as the values for the within-twin estimates. The estimated family effect, or the coefficients of the sum of the education of both twins, is larger than the return to education and significantly different from zero.

### 4.4. Measurement error

An important issue for twins studies is the measurement error problem. It is well known that classical errors in the measurement of schooling lead to a downward bias in the estimate of the effect of schooling on earnings, and that the fixed effects estimator magnifies such measurement error bias (Woodridge, 2002).

[10] These OLS estimates are very close to those obtained using the large NBS sample (Zhang et al., 2005).

[11] We run the same set of regressions using the sample of fraternal (DZ) twins. The results are reported in Table 6. The OLS estimate of the return to education for DZ twins is very similar to that for MZ twins, but the fixed effects estimate for DZ twins is larger than that for MZ twins, which suggests that there is some remaining bias in the within-twin estimates using DZ twins that can be removed in the within-twin estimates using the MZ twins sample. By taking estimates from the two samples together, we can draw the conclusion that one more year of schooling is associated with additional earnings of about 8.4% (see the OLS estimate in Table 3), of which 4.1% is due to the family effect (part of it is also due to the common genes of DZ twins), 1.4% is due to genetic effects, and 2.7% is the return to education.

One way to solve the problem of measurement error bias is to use the instrumental variable method. In this study, we follow the innovative approach of Ashenfelter and Krueger (1994) to obtain good instrumental variables. Specifically, in our survey we asked each twin to report both their own education and their co-twin's education. Where there is a risk of measurement error in the self-report of education, the cross-report is a potentially good instrument, as the report of the other twin should be correlated with the true educational level of a twin but uncorrelated with any measurement error that might be contained in the self-report.

Following Ashenfelter and Krueger (1994), the instrumental variable approach can be applied as follows. Writing $Z_j^k$ for twin $k$'s report of twin $j$'s schooling and assuming classical measurement error, we use $Z_1^1 - Z_2^2$ as the regressor and $Z_1^2 - Z_2^1$ as the instrumental variable (IV) in the differenced log earnings equation. This approach is valid even in the presence of common family-specific measurement error, because the family effect is eliminated through differencing. We call this instrumental variable model the IVFE-1.

Before reporting the IV estimates, it is worth examining the correlations between the education variables. The correlations between the self and co-twin reports of the education of the same twin, or $cov(Z_1^1, Z_1^2)$ and $cov(Z_2^2, Z_2^1)$, are 0.932 and 0.923 in our sample, compared to 0.920 and 0.877 in the sample of Ashenfelter and Krueger (1994). These higher correlations suggest that the measurement error is smaller than that in their sample, as these correlations are estimates of the reliability ratio of the education measures. The high correlations also suggest that the co-twin-reported level of education is a good instrumental variable for the self-reported level of education in our sample.

The IVFE-1 estimates that are reported in the first two columns of Table 4 show that measurement error has biased the fixed effects estimates in columns (5) and (6) of Table 3 downward, as in other studies in the literature. The IVFE-1 estimates of the return rise by about 22% (from 0.027 with the fixed effects model to 0.033 with the IVFE-1 model), which suggests that a fraction of the variability in the reported differences in years of schooling is due to measurement error.

However, the IVFE-1 estimates may also be biased if the measurement error terms in $Z_1^1 - Z_2^2$ and $Z_1^2 - Z_2^1$ are correlated. This will occur if there is an individual-specific component of the measurement error in reporting education. This motivates us to implement another instrumental variable that will be valid even in the presence of correlated measurement errors. To eliminate the individual-specific component of the measurement error in the estimation, it is sufficient to use $Z_1^1 - Z_2^1$ as the regressor and $Z_1^2 - Z_2^2$ as the IV (Ashenfelter and Krueger, 1994). We call this estimator IVFE-2. The new estimates of the return to education are 3.6–3.8% (last two columns of Table 4), which are about 15% greater than the IVFE-1 estimates. Overall, our IVFE estimates of the return to education are larger than the fixed effects estimates, which suggest that measurement error has indeed biased the within-twin estimates downward.

### 4.5. Potential biases of within-twin-pair estimates

Ashenfelter and Rouse (1998) emphasize that there can be no genetic differences between identical twins except by measurement error and argue that different schooling levels of identical twins are due to random deviations that are not related to the determinants of schooling choices. However, Bound and Solon (1999) examine the implications of the endogenous determination of which twin goes to school for longer, and conclude that twins-based estimation is vulnerable to the same sort of bias that affects conventional cross-sectional estimation. Bound and Solon (1999) and Neumark (1999) argue that although taking a within-twin-pair difference removes genetic variation, that is, it removes $\mu_i$ from Eq. (3), this difference may still reflect an ability bias to the extent that ability consists of more than just genes. In other words, within-twin-pair estimation may not completely eliminate the bias of conventional cross-sectional estimation, because the within-twin-pair difference in ability may remain in $\varepsilon_{1i} - \varepsilon_{2i}$ in Eq. (3), which may be correlated with $Z_{1i} - Z_{2i}$. If endogenous variation in education comprises as large a proportion of the remaining within-twin-pair variation as it does of the cross-sectional variation, then within-twin-pair estimation is subject to as large an endogeneity bias as cross-sectional estimation. We believe that the potential endogeneity of schooling differences between identical twins could correspond to remaining unobserved differences in ability or personality that may exist despite the common genetics or they may result from different experiences (early childhood illness, placement in different classes or schools, influence of different friends, and so on). The major concern of the within-twin-pair estimate is thus whether it is less biased than the cross-sectional estimate, and is therefore a better estimate.

Although within-twin-pair estimation cannot completely eliminate the bias of the cross-sectional estimator, it can tighten the upper bound on the return to education. As discussed, Ashenfelter and Rouse (1998), Bound and Solon (1999), and Neumark (1999) have debated the bias with cross-sectional and within-twin-pair estimation at length. Note that the bias in the cross-sectional estimator depends on the fraction of variance in education that is accounted for by variance in unobserved ability that may also affect earnings, that is, $\frac{cov(Z_i, \mu_i + \epsilon_i)}{var(Z_i)}$. Similarly, the ability bias of the fixed effects estimator depends on the fraction of within-twin-pair variance in education that is accounted for by within-twin-pair variance in unobserved ability that also affects earnings, that is, $\frac{cov(Z_i, \mu_i + \epsilon_i)}{var(Z_i)}$. If the endogenous variation within a family is smaller than the endogenous variation between families, then the fixed effects estimator is less biased than

**Table 4**
Instrumental variable fixed effects estimates of the return to education of Chinese MZ twins (dependent variable: log earnings).

| | IVFE-1 ($\Delta Z''$ as IV) | | IVFE-2 ($\Delta Z^{**}$ as IV) | |
|---|---|---|---|---|
| | (1) | (2) | (3) | (4) |
| Education ($\Delta Z'$) | 0.032* (0.019) | 0.033* (0.019) | | |
| Education ($\Delta Z^*$) | | | 0.036** (0.018) | 0.038** (0.018) |
| Married | | −0.043 (0.053) | | −0.048 (0.052) |
| Tenure | | 0.016 (0.010) | | 0.016 (0.010) |
| Twin pair | 488 | 488 | 488 | 488 |
| Observations | 976 | 976 | 976 | 976 |

Notes: $\Delta Z'$ is the difference between the self-reported education of twin 1 and the self-reported education of twin 2. $\Delta Z''$ is the difference between the education of twin 1 as reported by twin 2 and the education of twin 2 as reported by twin 1. $\Delta Z^*$ ($\Delta Z^{**}$) is the difference between twin 1's (twin 2's) report of his/her own education and his/her report of the other twin's education. Standard errors in parentheses are robust to heteroscadesticity.

* Significant at 10%.

** Significant at 5%.

**Table 5**
Between-families and within-twin-pair correlations of education and other variables (488 MZ twin pairs).

| Between-family correlations | Education | Within-twin-pair correlations | ΔEducation |
|---|---|---|---|
| Married | −0.1445*** (<0.01) | ΔMarried | −0.0173 (0.70) |
| Spousal education | 0.6172*** (<0.01) | ΔSpousal education | 0.1518** (0.02) |
| Party member | 0.2571*** (<0.01) | ΔParty member | 0.1166** (0.02) |
| Working in foreign firm dummy | 0.0904* (0.06) | ΔWorking in foreign firm dummy | 0.0214 (0.66) |
| Tenure | −0.2614*** (<0.01) | ΔTenure | −0.1253*** (0.01) |

Notes: The significance levels are in parentheses. The between-family correlations are the correlations between average family education (average of the twins) and average family characteristics, and the within-twin-pair correlations are the correlations between the within-twin-pair differences in education and the within-twin-pair differences in other characteristics.
* Significant at 10%.
** Significant at 5%.
*** Significant at 1%.

the cross-sectional estimator. Hence, even if there is an ability bias in the within-twin-pair regressions, the fixed effects estimator can still be regarded as a lower upper bound on the return to education if we are confident that education and ability are positively correlated. In that case, we can credit the within-twin-pair estimates with having tightened the upper bound on the return to education.

To examine whether the within-twin-pair estimate is less biased than the cross-sectional estimate, we follow Ashenfelter and Rouse (1998) and conduct a correlation analysis. We use the correlations of average family education over each twin pair with the average family characteristics that may be correlated with ability (for example, marital status, spousal education, membership of the Chinese Communist Party, working in a foreign firm, and job tenure) to indicate the expected ability bias in a cross-sectional OLS regression. We then use the correlations of the within-twin-pair differences in education with the within-twin-pair differences in these characteristics to indicate the expected ability bias in a within-twin-pair regression. If the correlations in the cross-sectional case are larger than those in the within-twin-pair case, then the ability bias in the cross-sectional regressions is likely to be larger than the bias in the within-twin-pair regressions.

The correlation tests that are reported in Table 5 suggest that the within-twin-pair estimation of the return to education may indeed be less affected by omitted variables than the cross-sectional OLS estimation. Note that the between-family correlations are all larger in magnitude than the within-twin-pair correlations. For example, the correlation between average family education and average spousal education is as large as 0.62 (column 1, row 2), which suggests that twins in families with a high average level of education marry highly educated spouses. This is consistent with the assumption that spousal education reflects an individual's ability and family background. The correlation of the within-twin-pair difference in education and the within-twin-pair difference in spousal education is about a quarter of the between-family correlation. This suggests that, to the extent that spousal education measures ability, the within-twin-pair difference in education is less affected by ability bias than the average family education. However, this within-twin-pair correlation is still statistically significant and large in magnitude, which suggests that within-twin-pair differencing cannot completely eliminate the ability bias that is embodied in the estimated coefficient on education. Thus, the within-twin pair estimation may only establish a lower upper bound for the estimated return to education than that from a cross-sectional OLS estimation. The correlations of education with some other variables provide similar evidence that the within-twin-pair estimation is subject to a smaller omitted ability bias. Of course, these characteristics are only an incomplete set of ability measures, but the evidence is suggestive.

As argued by Bound and Solon (1999), a key assumption for the conclusion that the FE estimate establishes a lower upper bound is that education is positively correlated with unobserved ability. To examine the sign of the correlation between education and unobserved ability, we compare the estimates using MZ twins to those using DZ twins. For DZ twins, the genes of the two twins are different, and thus there remain more unobserved ability in the error term in the DZ estimates compared to the MZ estimates. Thus, if DZ estimates are larger (smaller) than the MZ estimates, then it means that the unobserved ability is positively (negatively) correlated with education.

Comparing DZ estimates to MZ estimates indeed suggests that education is positively correlated with unobserved ability. Regression results of DZ estimates are reported in Table 6, while those for MZ twins are reported in Tables 2 and 3. Note that the OLS estimate of the return to education (8.5%) using the DZ sample is almost the same as the OLS estimate using the MZ sample (8.4% in column 4 in Table 3). However, the FE and IVFE estimates for the DZ sample are much larger than those for the MZ sample, suggesting that the (additional)

**Table 6**
OLS and fixed effects estimates of the return to education for DZ twins in urban China (dependent variable: log earnings).

| Sample model | DZ twins | | | |
|---|---|---|---|---|
| | OLS (1) | FE (2) | IVFE-1 (ΔZ″ as IV) (3) | IVFE-2 (ΔZ** as IV) (4) |
| Own education | 0.085*** (0.008) | 0.044** (0.012) | 0.054*** (0.017) | 0.056*** (0.018) |
| Age | 0.067** (0.028) | | | |
| Age squared | −0.105*** (0.038) | | | |
| Gender (male) | 0.164*** (0.054) | | | |
| Married | 0.057 (0.055) | 0.085 (0.065) | 0.089 (0.065) | 0.103 (0.064) |
| Tenure | 0.010 (0.009) | −0.021 (0.018) | −0.021 (0.018) | −0.022 (0.018) |
| Twin pairs | 316 | 316 | 316 | |
| Observations | 632 | 632 | 632 | 632 |
| R-squared | 0.19 | 0.05 | | |

Notes: As in Table 4, for Model IVFE-1, ΔZ′ is the difference between the self-reported education of twin 1 and the self-reported education of twin 2. ΔZ″ is the difference between the education of twin 1 as reported by twin 2 and the education of twin 2 as reported by twin 1. For Model IVFE-2, ΔZ* (ΔZ**) is the difference between twin 1's (twin 2's) report of his/her own education and his/her report of the other twin's education. Standard errors in parentheses are robust to heteroscadesticity and clustering at the family level.
** Significant at 5%.
*** Significant at 1%.

unobserved ability in the DZ sample is positively correlated with education.

One potential endogeneity that could be more problematic is if the more educated twin helps the less educated twin in some way (e.g. finding a better job). We tested this possibility by regressing own earnings on the education level of the more educated twin (in addition to controlling for own education) and found no significant effect.

Finally, if one twin stayed in another city, the twin pair would not be in our sample. "Splitting" twins would be a problem, but there is little we could do about it. To the extent that the more able one is more likely to migrate to another city, our within-twin estimate based on a sample in which both twins were not separated into two cities would tend to under-estimate the schooling return. However, according to the 1% sample of China's 2000 census, we find that the inter-urban migration rate was only 0.975%, and thus any sample selection bias arising from splitting twins should not be a major issue.

## 5. Why is return in urban China low?

It is interesting to compare our estimates to other estimates in the literature that draw on data from different countries, namely, the US, UK and Australia. Note first that our estimate of the raw return to education, that is, the cross-sectional OLS estimate, is 8.4% (column 4 of Table 3), which is very close to other estimates in the literature (column A of Table A1). However, our within-twin-pair estimate is only 2.7%, which is smaller than most estimates in the literature. Moreover, the ability bias in our sample, which stands at 5.7%, is larger than the ability bias that has been found in all other studies (column C of Table A1). These results together suggest that, in our sample drawn from urban China, ability bias is relatively large compared to other countries.

To ascertain why the true return to education is so low and the ability bias so high in urban China, we need to understand the Chinese education system, because this education system may explain the high ability bias and low return to education in these estimates.

### 5.1. The Chinese education system

The Chinese education system is highly selective and exam oriented (Zhang, 2009). It is composed of two stages: the compulsory stage and the non-compulsory stage. The compulsory stage comprises six years of primary school and three years of junior high school.[12] Currently, most urban children finish nine years of compulsory education. Junior high school graduates have a choice of attending high school or vocational high school,[13] and are required to take an entrance exam to gain a place at either type of institution. High school graduates are eligible for the college entrance exam, but vocational school graduates are generally not. In our monozygotic (MZ) twins sample, 74% had a high school or vocational school degree or above.[14] Because of the huge number of people aspiring for higher education and the limited number of places at colleges and universities, entrance to college is extremely competitive. Only 13% of the workers in our sample obtained a college degree. To select those who will go on to college education, a nationwide college entrance exam system has been adopted, and the exam days of June 7, 8, and 9 determine the future of many young people each year.[15] Those who pass the examinations will become white collar workers, and those who fail will most likely become blue collar workers.

Because of the competitive nature of the education system, schools, and in particular junior high schools and high schools, place great emphasis on exam-taking techniques.[16] Although high school in China lasts for three years, the whole curriculum is normally finished in one and a half years or even a shorter time, with the rest of the time being spent on preparation for the college entrance exams. Although the first half of high school teaches students new things, the teaching is also focused on exam-type problem-solving techniques. High school students need to finish a lot of homework every day, and normally need to go to school on weekends and vacations. All of this extra time is spent on training students to solve exam questions. Schools and teachers are rewarded largely on the basis of the success rate of their students in the entrance exams, and thus have no incentive to teach them anything else. These exam-taking techniques very often have little to do with the knowledge and skills that are needed for life and work (Han and Yang, 2001), and it is thus unsurprising that such kind of schooling has a low return in the workplace. Using a large cross-sectional urban household data set, Zhang et al. (2005) find that returns to technical school are much higher than those to high school. This indicates that technical schools do teach students productive vocational skills whereas high schools enhance skills that are not as productive in the workplace. Indeed, both vocational high schools and colleges are much less exam oriented than high schools.[17]

Two other features of the Chinese non-tertiary education system may also damp the return. First, the curricula (*jiao xue da gang* in Chinese) for primary school, junior high school, and high school are fixed by the Ministry of Education, and the most important part of these curricula is to specify what should be covered by the high school and college entrance exams. Second, high school students have to decide to take either arts or science for the rest of their education. Both arts and science students take Chinese, English, and political science, but arts students take geography, history, and basic mathematics, whereas science students take physics, chemistry, biology, and advanced mathematics.

These features of the Chinese education system can help to explain why the omitted ability bias (or selection effect) is high and the true return to education low in our estimations. Because of intense competition, only the very talented can advance to higher education, and thus education (or entrance exams) is a very good selection mechanism. Because of the exam-oriented education system, non-tertiary education, and in particular high school education, add little value to students in terms of general knowledge or workplace skills, except as a means of selecting talented candidates into college. High school graduates who are not able to get into college may thus have wasted three years on training in exam-taking techniques.

[12] There have been increases in the number of years in each level of non-tertiary schooling. In the 1970s, primary school was only 5 years, and junior high school and high school was 2 years at each level. In the 1980s, the old system has been transferred to the current system gradually. In different cities, the changes happened in different years. Some people in our sample went to school in the old system. In our sample, primary school took 5.7 years on average, junior high school took 2.9 years, and high school took 2.6 years.

[13] There are several types of vocational schools in China, which are called vocational schools, vocational high schools, or skilled workers' schools. In this paper, we group them and 2-year vocational college together under the term "vocational high schools."

[14] The percentage for the whole sample is 72%.

[15] The exam dates were formerly 7, 8, and 9 July, but were changed in 2003 to avoid the hot weather.

[16] It is no secret that the Chinese have very good exam taking skills. For example, among graduate school applicants in the United States, those from China normally have very high scores in GRE and other standard tests, and sometimes even have higher test scores in verbal English than native speakers. However, most Chinese people have never spoken English before coming to the United States because oral English is neither required by most US graduate schools nor emphasized in the English exams in China.

[17] First, vocational high schools or colleges have the freedom to choose their own curricula. Second, and most importantly, vocational schools and colleges are usually the final stage of education, and thus exams are no longer important.

### 5.2. What levels of education pay?

The exam-oriented education system not only helps to explain why the return is low and ability bias high, but also suggests that the return to education may differ across education levels. It seems that exam-oriented high school education is the least useful level of education, and is valuable only as a selection mechanism for colleges. This means that the education of high school graduates who do not make it to college should be least rewarded by employers. We investigate whether this is true by estimating the return to different levels of education.

In the literature on twins studies, years of schooling is generally used as the measure of education (see, for example, Ashenfelter and Krueger, 1994; Bonjour et al., 2003). However, little work has been carried out to examine the returns to different levels of education. As most of the literature draws on data from developed countries, and a large proportion of workers in these countries have some years of college education and have at least completed a high school education, it may not be necessary to examine the return to high school education versus the return to college education. However, knowing the returns to different levels of education is still very important for a developing country such as China, where college education remains very limited. Knowing the returns to different levels of education could help the government to better allocate limited resources for education.

We investigate whether high school education has a lower return than vocational school or college education by estimating the return to different levels of education. More specifically, we use four education dummies as measures of education, namely, high school, vocational high school, vocational college, and college dummies (junior high school is taken as the base group). The high school dummy equals 1 if the last qualification that an individual obtained was a high school qualification, and 0 otherwise, and other dummy variables are similarly defined. As described in Section 3, the within-twin variation in education levels is sufficiently large so that the performance of within-twin estimation should be good.

One concern is that the measurement error problem becomes much more involved when we use dummy variables to measure educational attainment, because the measurement error in dummy variables is not classical. To the best of our knowledge, there is no method that we could directly draw upon to correct the measurement error in the dummy variables in our twins estimations. However, Black et al. (2000) provide a method that establishes the lower and upper bounds of the true return to education as estimated using a twins sample. According to them, the IVFE-1 estimates establish an upper bound, and the lower bound is set by fixed effects estimates using the sample of twins in which both twins agree on the other's educational level. This lower bound should be tighter (larger) than the fixed effects estimate using the whole twins sample.

It is also worthwhile investigating the size of the potential measurement error by comparing the self-reported education and the cross-reported education as stated by the co-twin. Although the reports of 10.5% of the twins do not agree on the number of years of education, only 2.3% of them do not agree on the level of education, which translates to 19 pairs of twins (out of 488 pairs) disagreeing on the within-twin difference in educational level. The low potential reporting error for educational level simply reflects the fact that reporting the final degree requires less information and calculation than reporting the number of years of schooling. The low percentage of disagreement over the level of education suggests that the bias that is caused by measurement error may not be too great in our estimations. Indeed, the reliability ratios for these education dummies are as high as 0.96–0.98, which suggests that only about 2–4% of the measured variance in these education dummies is due to measurement error.

The regression results that are reported in Table 7 show that the return to high school education is much lower than the return to vocational school and college education. The high school dummy is positive and significant in the OLS estimation (column 1), but the return to high school education is much lower than the return to vocational high school, vocational college or college education, and the differences are significantly different from zero. As both high school and vocational high school usually require three additional years of education beyond junior high school, their returns are directly comparable. The large difference between the two returns suggests that it pays more to attend vocational high school, relative to academic high school only.

The large difference between the return to high school education and the return to vocational high school, vocational college or college education remains for the fixed effects and IVFE estimates.

**Table 7**
Various estimates of the return to different levels of education for MZ twins in urban China.

| Sample | Dependent variable: log earnings | | | | | |
|---|---|---|---|---|---|---|
| | MZ twins | MZ twins | MZ twins (restricted sample) | MZ twins | MZ twins (restricted sample) | MZ twins |
| Model | OLS | FE | FE | IVFE-1 | FE | IVFE-1 |
| | (1) | (2) | (3) | (4) | (5) | (6) |
| High school | 0.130** (0.053) | 0.022 (0.078) | 0.040 (0.086) | 0.054 (0.086) | 0.020 (0.078) | 0.052 (0.086) |
| Vocational high school | 0.229*** (0.063) | 0.165* (0.085) | 0.196** (0.090) | 0.219** (0.093) | 0.154* (0.086) | 0.208** (0.094) |
| Vocational college | 0.493*** (0.051) | 0.190** (0.084) | 0.215** (0.091) | 0.230** (0.090) | 0.190** (0.085) | 0.230** (0.090) |
| College | 0.706*** (0.058) | 0.314*** (0.118) | 0.357*** (0.125) | 0.400*** (0.130) | 0.318*** (0.119) | 0.404*** (0.131) |
| Age | 0.045** (0.019) | | | | | |
| Age squared | −0.062*** (0.023) | | | | | |
| Gender (male) | 0.192*** (0.037) | | | | | |
| Married | −0.069 (0.050) | −0.032 (0.053) | −0.038 (0.054) | −0.029 (0.053) | −0.033 (0.053) | −0.030 (0.053) |
| Tenure | 0.008 (0.006) | 0.015 (0.010) | 0.015 (0.010) | 0.016* (0.010) | 0.015 (0.010) | 0.015 (0.010) |
| Send-down years | | | | | 0.032** (0.013) | 0.031** (0.013) |
| Twin pairs | 488 | 469 | 488 | 488 | 488 | |
| Observations | 976 | 976 | 938 | 976 | 976 | 976 |
| R-squared | 0.22 | 0.03 | 0.04 | 0.04 | | |

Notes: The regressions in columns (3) and (5) use a restricted sample with both twins agreeing on the within-twin difference of education levels. For Model IVFE-1, we use $\Delta Z'$ as independent variables, which are instrumented by $\Delta Z''$, where $\Delta Z'$ is the difference between the self-reported education of twin 1 and the self-reported education of twin 2 and $\Delta Z''$ is the difference between the education of twin 1 as reported by twin 2 and the education of twin 2 as reported by twin 1. Standard errors in parentheses are robust to heteroscadesticity and for OLS clustering at the family level.

* Significant at 10%.

** Significant at 5%.

*** Significant at 1%.

When we run the fixed effects estimation using the restricted sample (in which both twins agree on the within-twin difference in educational level) as the lower bound and the IVFE estimates as the upper bound, the return to high school education is bounded between 4.0 and 5.4% (columns 3 and 4 of Table 7). In contrast, the return to vocational high school education is bounded between 19.6 and 21.9%, the return to vocational college education is bounded between 21.5 and 23.0%, and the return to college education is bounded between 35.7 and 40.0%. The return to college education is similar to that in the US and Europe, where the estimate of each additional year of education is about 10%. Moreover, for each of the regressions (columns 2–4), the estimated coefficient of the high school dummy is significantly different from that of the vocational school and college dummies. The lower return to high school education provides an explanation for why the overall return to education is low in China.

As many workers in our sample were sent down to the countryside during the Cultural Revolution and their education was negatively affected by this movement (Li et al., 2010), one may wonder whether send-down affects the estimates of the return to education. To address this concern, we report regressions that also include the send-down years as a control variable in the last two columns of Table 7. Similar to findings in Li et al. (2010), we find that send-down has a positive impact on earnings, and controlling for send-down years does not affect our estimates to the returns to different levels of education.

As discussed earlier, to examine the sensitivity of the results to potential sample selection problems, we enlarged the sample by including twins who were not working at the survey time and using their earnings of the last jobs (allowing for age and inflation adjustment). The results reported in Table A2 indicate qualitatively the same message as before, that is, there is no return to high school per se but a large return to vocational or college education.

## 6. Conclusion

In this paper, we empirically measure the return to education in urban China. By using twins data to control for omitted ability and family effects, we find that most of the return to education that is estimated by the OLS model is actually due to the effects of unobserved ability or family background. In other words, our findings suggest that the selection role of education is more important than the knowledge that is acquired at school. The fixed effects estimate of the return to one year of education is only 2.7%, and the part of the education premium that is due to unobservable ability is as large as 5.7%. Our sensitivity analyses suggest that the within-twin estimates are biased downward due to measurement error, and when we subsequently correct this bias the estimated return rises to 3.8%. We further show that the return to high school education is almost zero, whereas the returns to vocational high school, vocational college and college education are 19.6–21.9%, 21.5–23.0% and 35.7–40.0%, respectively. These results suggest that the low overall return to years of schooling is caused by the low return to high school education.

The earlier findings of a low return to education in China, such as those of Byron and Manaloto (1990) and Meng and Kidd (1997), have generated a great deal of interest among economists. In the search for explanations of the low return, most economists have turned to the remaining elements of the planned economy. We agree that the return was low when the Chinese economy was under a planned regime and is likely to rise as the economy becomes more market-oriented (see, for example, Heckman and Li, 2004; Zhang et al., 2005), but argue that economic transition may not be the whole story. In fact, the returns to the various levels of education are different, and in particular we find that exam-oriented high school education has no return in terms of earnings; rather, it may merely serve as a selection mechanism for colleges. Few previous studies have paid attention to the exam-oriented education system in China (or that of other countries) and its consequences for the return to education.

Knowing the true return to education has important policy implications. Our findings show that the return to education is not universally low in China. The return to vocational school or college is comparable to estimates that draw on twins data from other countries such as the US and UK. Thus, our findings support the argument of Heckman and Li (2004) and Fleisher and Wang (2004) that investing in human capital is worthwhile in China. Furthermore, our finding that the return to high school education is zero suggests that the government should carefully review the high school education in terms of its curriculums and examinations.

## Appendix

**Table A1**
Estimated return to years of education using different twins samples.

| Study | Sample and country | OLS (A) | FE (B) | Omitted variable bias (C = A − B) | IVFE (D) |
|---|---|---|---|---|---|
| Taubman (1976a) | NAS-NRC Twin Registry sample of white male army veterans, USA | 0.079 | 0.027 | 0.052 | – |
| Ashenfelter and Krueger (1994) | Twinsburg sample, USA | 0.084 | 0.092 | −0.008 | 0.129 |
| Behrman et al. (1994) | NAS-NRC Twin Registry, Minnesota Twin Registry, USA | – | 0.035 | – | 0.050 |
| Miller et al. (1995) | Australia Twin Registry | 0.064 | 0.025 | 0.039 | 0.048 |
| Behrman et al. (1996) | Female twins born in Minnesota, USA | – | 0.075 | – | – |
| Ashenfelter and Rouse (1998) | Twinsburg sample, USA | 0.110 | 0.070 | 0.040 | 0.088 |
| Behrman and Rosenzweig (1999) | Minnesota Twin Registry, USA | – | – | – | 0.104 |
| Bonjour et al. (2003) | Twins Research Unit, St., Thomas' Hospital (female only), London, UK | 0.077 | 0.039 | 0.038 | 0.077 |

**Table A2**
Various estimates of the return to different levels of education for MZ twins in urban China (using previous job's earnings if current earnings are missing).

| | Dependent variable: log earnings | | |
|---|---|---|---|
| Sample | Twins | Twins (restricted sample) | Twins |
| Model | FE | FE | IVFE-1 |
| | (1) | (2) | (3) |
| High school | 0.021 (0.091) | 0.040 (0.099) | 0.120 (0.112) |
| Vocational school | 0.251*** (0.091) | 0.281*** (0.097) | 0.316*** (0.104) |
| College | 0.365*** (0.122) | 0.406*** (0.129) | 0.517*** (0.129) |
| Married | −0.042 (0.055) | −0.045 (0.057) | −0.040 (0.056) |
| Tenure | 0.029** (0.014) | 0.030** (0.014) | 0.031** (0.014) |
| Twin pairs | 611 | 586 | 611 |
| Observations | 1222 | 1172 | 1222 |
| R-squared | 0.03 | 0.04 | |

Notes: The regression in column (2) uses a restricted sample with both twins agreeing on the within-twin difference of education levels. For Model IVFE-1, we use $\Delta Z'$ as independent variables, which are instrumented by $\Delta Z''$, where $\Delta Z'$ is the difference between the self-reported education of twin 1 and the self-reported education of twin 2 and $\Delta Z''$ is the difference between the education of twin 1 as reported by twin 2 and the education of twin 2 as reported by twin 1. Standard errors in parentheses are robust to heteroscadesticity. * significant at 10%; ** significant at 5%; *** significant at 1%.

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