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Education expansion and returns to schooling in urban China, 2001−2010: evidence from three waves of the China Urban Labor Survey

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This study examines the effect of the expansion in education that occurred in the first decade of the twenty-first century on the returns to schooling in urban China for migrants and non-migrants using three waves of the China Urban Labor Survey (CULS), corresponding to 2001, 2005 and 2010. Our main finding is that the premium to education increased by about 2%−3% over a period in which there was a rapid increase in education levels. This result is consistent with the demand for skilled labor increasing at a time when China tries to move up the value-added chain and an observed increase in urban wage inequality. We find that the education premium is higher for non-migrants than migrants and higher for males than females.

Keywords: China; education expansion; returns to schooling

Introduction

China underwent rapid expansion in post-secondary education in the first decade of the twenty-first century. In 1999 the Chinese government decided to increase post-secondary enrolments by 40% in response to weak aggregate demand and deflation in the late 1990s. The number of post-secondary graduates tripled between 2002, which was when the first group of post-secondary students with three-year degrees graduated following the 1999 reforms, and 2010 (Hu and Hibel, forthcoming). The expansion in Chinese higher education in the first decade of the new century has been described as representing the transition from 'elite' to 'mass education' (Wu 2014).

There were two reasons for increasing post-secondary enrolments in the late 1990s. The first was a hope that such a step would stimulate aggregate demand, when accompanied by an increase in tuition fees, in a country that values education. The second reason was that it was hoped that this move would increase the proportion of skilled labor in the workforce needed to stimulate economic growth and meet the challenges of future growth in the new century (Meng, Shen, and Sen 2013).

The main outcome of the transition from elite to mass education was a substantial increase in the average years of schooling after 2002. Based on the China Urban Labor Survey (CULS) data, which will be employed in this paper, average years of schooling for the urban population as a whole increased from 10.39 years in 2001 to 10.44 years in 2005 to 11.46 years in 2010. Based on the CULS data, there was also an increase in average years of schooling for both non-migrants and migrants.[1] For non-migrants, average years of schooling increased from 11.89 years in 2001 to 12.01 years in 2005 and

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12.61 years in 2010. For migrants, the corresponding increase was from 8.6 years in 2001 to 8.96 years in 2005 to 10.33 years in 2010.

The effect of an expansion in higher education, and resultant increase in average years of schooling, on the returns to schooling is *ex ante* unclear. On the one hand, on the supply side one would expect education expansion to lower the returns to schooling if the increasing supply of post-secondary graduates results in 'over-education' (Wu 2014). In this respect, there is evidence that the expansion of higher education in China is responsible for an increase in unemployment among college graduates (Li, Whalley, and Xing, forthcoming). As documented by Freeman (2009) and Li, Whalley, and Xing (forthcoming), an increase in the female education level is a prominent feature of higher education expansion. An increase in the participation rate of women, in particular educated women, is another supply side factor that could result in a decrease in the returns to schooling. The reason is that such a development increases the supply of educated professionals, reducing the wage employers' need to pay to attract such individuals. On the other hand, demand side factors, such as skill-biased technological change and workplace innovations, can be expected to reward those with higher education and increase the returns to schooling. China has experienced increased urban income inequality throughout the first decade of the twenty-first century (see e.g. Meng, Shen, and Sen 2013), which would be consistent with an increase in returns to schooling.

The issue of whether an expansion in higher education results in an increase, or decrease, in returns to schooling has important public policy implications. Expanding higher education has been seen in China (Wang 2013) and other countries, such as the United States and the United Kingdom (Hu and Hibel, forthcoming), as a means of increasing living standards and improving social equity. As Ashenfelter and Rouse (2000, 111) put it: "school is a promising place to increase the skills and incomes of individuals. As a result, educational policies have the potential to decrease existing, and growing, inequalities in income". In the Chinese context, Heckman (2005) has suggested that "human capital is the asset that ultimately determines the wealth of China. Fostering access to education will reduce inequality in the long run".

However, to the extent that educational expansion is not equitable and demand side factors dominate, productivity increases may occur at the expense of social equity and education expansion might actually exacerbate economic inequality, rather than promote social equity (Lemieux 2006). There is evidence that this reflects what has happened in China. Since the mid-1990s, the deepening of economic reform in urban areas has enhanced flexibility and diminished administrative control in the workplace. As a result, wages in China have become increasingly more responsive to productivity and wage differentials have increased. Throughout the 2000s, a growing proportion of post-secondary students were forced to fund their own educational expenses (Heckman 2005) and there was a rapid increase in tuition fees (Wang 2013).

The purpose of this paper is to examine whether the education expansion, and commensurate increase in average years of schooling, has 'devalued' the economic returns to education in the Chinese labor market throughout the 2000s. To do so, we estimate the private returns to education using three waves of the CULS for 2001, 2005 and 2010. We extend the existing literature for China in at least three ways.

First, we use a representative survey, administered in a consistent manner, to examine changes in returns to schooling over the first decade of the twenty-first century. Most studies of the returns to schooling in urban areas use a single year of data. Differences in the peculiarities of datasets and differences in the methodologies and specifications make it difficult to compare results across time. Some studies have examined trends in returns to schooling using a consistent dataset, but these mainly focus on the 1980s and 1990s (see e.g. Appleton, Song, and Xia 2005; Qiu and Hudson 2010; Zhang et al. 2005).

Second, methodologically, we employ a twofold strategy to address the endogeneity of years of schooling. We use spouse's education as a conventional instrumental variable (IV) as well as a novel identification strategy, proposed by Lewbel (2012), which utilizes a heteroskedastic covariance restriction to construct an internal IV. For a long time, ordinary least squares (OLS) were used to estimate returns to schooling in China, which reflected the lack of suitable IVs in commonly used datasets. Thus, as recently as 2005, Li et al. (2005) stated: "Despite the rapid accumulation of evidence on the returns to education in China, no study has yet established causality". More recently, a series of studies have used various conventional IVs to identify the causal effect of education on earnings in urban China (see e.g. Heckman and Li 2004; Fleisher et al. 2004; Li and Luo 2004; Chen and Hamori 2009; Fang et al. 2012; Kang and Peng 2012; Mishra and Smyth 2013).

Common IVs have been parents' education, family background characteristics and spouse's education. However, each of these IVs has been criticized for not satisfying the exclusion restriction. Given these concerns, using the Lewbel (2012) method, which does not rely on satisfying the exclusion restriction in addition to a conventional IV, provides the results with added robustness. If the results using these identification strategies are similar, this should increase confidence in the findings.

The third contribution is that we explicitly consider returns to schooling for both migrant and non-migrant samples. An advantage of CULS is that over each of its three waves it contains data on migrants and non-migrants living in urban areas. There are very few datasets with representative data on education and wages of rural–urban migrants in China (Liu and Zhang 2012). Datasets employed in previous studies that have examined trends in returns to schooling over time (e.g. the urban household surveys employed by Zhang et al. 2005) do not contain data on rural–urban migrants.[2] Yet, excluding rural–urban migrants excludes an important segment of the urban labor market and this is increasingly the case since 2000. In 2011, there were approximately 160 million rural–urban migrants in the urban areas of China, constituting 44% of the urban labor force (National Bureau of Statistics of China (NBSC) 2012). In some developed coastal areas, such as Fujian and Shanghai, this figure is likely to be in excess of 50% (Gagnon, Xenogiani, and Xing 2011).

Returns to education in urban China

Table 1 provides an overview of a non-exhaustive list of existing studies. In the 1980s and early 1990s, returns to schooling in the Chinese urban labor market were generally low. In three of the earlier studies, Byron and Manaloto (1990), Johnson and Chow (1997) and Liu (1998) reported a rate of return in the range of 3%–4%. Li (2003) was the first to control for heterogeneity in working hours. He found a return of 5.4%. In the mid-1990s, returns to schooling in urban China were slightly higher, in the range of 5%–6% (see e.g. Bishop and Chiou 2004), but were still low compared with the world average of 10.1% (Psacharopoulos 1994).

Returns to schooling have increased since the mid-1990s. Ge and Yang (2011) found that the rate of return to one additional year of schooling in urban China, estimated using OLS, increased from 3.6% in 1988 to 11.4% in 2007. The figure from Ge and Yang (2011) for 2007 is consistent with OLS estimates for urban China for 2005 reported in Qian and Smyth (2008) of 12%–13%. Other studies report slightly lower estimates using OLS. Qiu and Hudson (2010) found that returns to schooling in urban China at the

Table 1. Selected existing studies of returns to schooling in Urban China.

Study	Sample	Years covered	Method	Estimate (%)
Byron and Manaloto (1990)	Urban China	1986	OLS	1.4
Johnson and Chow (1997)	Urban China	1988	OLS	3.3
Liu (1998)	Urban China	1988	OLS	2.9–3.6
Li (2003)	Urban China	1995	OLS	4.7–5.4
Bishop & Chiou (2004)	Urban China	1988, 1995	OLS	2.8, 5.6
Fleisher et al. (2004)	Urban China	1988–2002	IV	16.9–38.6
Heckman and Li (2004)	Urban China	2000	OLS, IV	7.3–23.2
Li and Luo (2004)	Urban China	1996	IV	15
Zhang et al (2007)	Urban China	2002	Twins	3.8–9.8
Giles, Park, and Wang (2008)	Urban China	2000	IV	8.3–9.6
Qian and Smyth (2008)	Urban China	2005	OLS	12–13
Chen and Hamori (2009)	Urban China	2004, 2006	OLS, IV	OLS: 7.7–8.1; IV: 12.6–14.5
Qiu and Hudson (2010)	Urban China	1989, 1993, 1997, 2000	OLS	5.1–6.9
Ge andYang (2011)	Urban China	1988–2011	OLS	3.6–11.4
Fang et al. (2012)	Urban China	1997–2006	IV	20
Kang and Peng (2012)	Urban China	1989–2009	OLS, IV	OLS: 2.2–10.3; IV: 6.2–11.0
Liu & Zhang (2012)	Urban China	1991–2009	OLS	0.5–7.4
Mishra and Smyth (2012)	Shanghai	2007	Lewbel (2012) IV	25–30
Ren and Miller (2012)	Urban China	2004, 2006	OLS	7–8
Mishra and Smyth (2013)	Ethnic Koreans in urban China	2009–2010	IV	21–23
Wang (2012)	Urban China	1995, 2002	IV	9.5–44
Wang (2013)	Urban China	1995, 2002	OLS, IV	OLS: 3.6–8.1; IV: 4.4–11.8
Mishra and Smyth (2014)	Shanghai	2007	OLS	6.9–7.4

beginning of the 2000s were 7%. Mishra and Smyth (2014) report that returns to schooling in Shanghai in 2007 were 6.9%–7.4%. Similarly, Chen and Hamori (2009) and Ren and Miller (2012) found returns to education in urban China in 2004–2006 were in the range of 7%–8%.

The labor planning system in the pre-reform (late 1950s to late 1970s) and early marketization periods contributed to the unequal distribution of income. Labor bureaus allocated labor and wages were largely fixed and depended on seniority. A further factor was that there was virtually no labor market. The state-owned sector dominated industry and dominated the allocation of labor and wage setting.

This started to change through the 1980s and 1990s in response to government initiatives to improve incentives and boost productivity. Specific policies of note included the introduction of a system of labor contracts in 1986, which was a forerunner for more flexibility in the labor market, and the 1994 Labor Law, which built a tighter connection between employee effort and monetary reward.

The expansion in the higher education sector, coupled with the increased cost of education, has focused attention on the size of the college premium. Studies find a sizeable college premium, reflecting the traditional shortage of college graduates in urban China (Fleisher et al. 2004; Heckman and Li 2004; Wang 2012).

Since the mid-2000s, some studies have used an IV approach (Chen and Hamori 2009; Fang et al. 2012; Fleisher et al. 2004; Heckman and Li 2004; Kang and Peng 2012; Li and Luo 2004; Mishra and Smyth 2012, 2013; Zhang, Liu, and Yung 2007). Most of these studies use conventional IVs for own education. Mishra and Smyth (2012) and Wang (2012) exploit heteroskedasticity in the data as an identification strategy. Mishra and Smyth (2013) use a combination of conventional and internal IVs. Fang et al. (2012) use the introduction of the Compulsory Education Law in 1986 as a natural experiment to address endogeneity of education. Giles, Park, and Wang (2008) use the differential impact of the Cultural Revolution across time and geographical location as an IV for schooling. As in the literature more generally, overall studies using an IV approach have found a higher rate of return than studies that have used OLS. In a meta-analysis of estimates of returns to schooling in China, Liu and Zhang (2012) found that estimates using IVs were 3.2%−3.6% higher than OLS estimates.

There is considerable evidence of a growing gender wage gap in urban China for both the urban population as a whole and for rural−urban migrants. Moreover, existing studies suggest that a significant portion of the wage gap cannot be explained by differences in human capital and other observed characteristics (see e.g. Gustafsson and Li 2000; Zhang et al. 2008; Magnani and Zhu 2012). Traditionally most studies have found that returns to schooling are higher for females than for males. Deolalikar (1993) argued that males have a comparative advantage in physical strength so that schooling becomes relatively more important to females, whose comparative advantage is in skill-intensive jobs. Li (2003) suggested that the higher returns to schooling for females reflect the relative dearth of highly educated females in urban China. Several recent studies have found that returns to schooling are higher for males than for females since 2000 (Chen and Hamori 2009; Qian and Smyth 2008; Kang and Peng 2012; Ren and Miller 2012). Some studies have found that OLS estimates are biased downward for males and biased upward for females (Chen and Hamori 2009; Kang and Peng 2012).

Data

We use household data from the three waves of the CULS (2001, 2005 and 2010). The CULS was designed by an international research team from China, the United States, the United Kingdom and Australia. It was administered by the Institute of Population and Labor Economics at the Chinese Academy of Social Sciences, working in collaboration with local National Bureau of Statistics Survey teams. In each wave, the CULS was administered to samples of migrant and non-migrant households.

The 2001 survey was administered in five provincial capital cities (Shanghai, Wuhan, Shenyang, Fuzhou and Xi'an). For the non-migrant household sample, proportional population sampling was used to sample an average of 10 households in each of 70 neighborhood clusters, where the clusters were selected using the 2000 Chinese Population Census. For the migrant sample, 60 communities were selected using the 2000 Chinese Population Census. On average, 10 non-migrant households were interviewed in each neighborhood cluster and 10 migrants were interviewed in each community in each city. Thus, in total, in each city, all individuals aged 16 years and above in 700 non-migrant households and 600 individual migrants were surveyed.

The 2005 survey was administered in the same five provincial capital cities, but, in addition, seven municipal cities (Wuxi, Yichang, Benxi, Zhuhai, Shenzhen, Baoji and Daqing) were added. Using the same sampling approach as the 2001 survey, 500 non-migrant households and 500 migrant households were investigated in each of the five provincial cities, and 400 migrant households were investigated in each of the seven municipal cities. In each household, all family members who were aged 16 years or above and who were no longer in school were individually interviewed.

Following the same sampling method as the 2001 and 2005 surveys, the 2010 survey was carried out in the same five provincial capital cities covered in the earlier surveys, plus Guangzhou. The seven municipal cities included in the 2005 survey were not included in the 2010 survey. In the 2010 survey, the survey was administered to 700 non-migrant households and 600 migrant households in each of the six cities.

To ensure the results are comparable across time, this study only uses data from the five provincial capital cities surveyed in all three years; that is, Shanghai, Wuhan, Shenyang, Fuzhou and Xi'an. Given our focus is on returns to schooling, we restrict the sample to urban non-agricultural workers aged 16−64 years who are in fulltime employment, and who have complete information on education and earnings. In the aftermath of the Fifteenth Party Congress in 1997 there were large-scale numbers of workers laid-off from China's state sector. In 1997−1998, it is estimated that around 12 million workers in the state-owned sector were laid-off (Meng 2000). Dong and Xu (2008) show that labor market restructuring in the state-owned sector continued into the 2000s. Usually these workers retained their ties with the enterprise, and thus were not officially unemployed, but were typically only paid base wages and did not report for work each day. We exclude such workers from the analysis.

A common approach in the literature is to exclude the self-employed from the sample. In the results below we include the self-employed, given the rise in the prevalence of self-employment in urban China throughout the 2000s, reflecting the influx of rural−urban migrants into Chinese cities. A high proportion of rural−urban migrants are self-employed in Chinese urban areas, typically averaging 50%−60% (Wang, Cai, Zhang 2010). In each wave of the CULS, the self-employed constitute in excess of one quarter of the sample, with the proportion of self-employed migrants much higher than non-migrants. Arabsheibani and Mussurov (2007) adopt a similar approach in their study of the returns to schooling in Kazakhstan, where the level of self-employment is also high. Similarly, Mishra and Smyth (2013) in a study of returns to schooling for a sample of ethnic Koreans from China's northeast follow this strategy, given that a high proportion of this minority group are self-employed in China's urban labor market.[3] The numbers of respondents for which we had valid data were 6197 in 2001, 6490 in 2005 and 8956 in 2010.

Table 2 provides descriptive statistics for the full sample and for migrant and non-migrant respondents in each of the three years studied. In 2001, approximately 50% of the sample worked in private enterprises or individual businesses. This figure increased to approximately 70% in 2010. The proportion of migrants working in private enterprises and individual businesses is considerably higher than non-migrants in each of the three years. In 2001, 28% of the sample worked in collective or state-owned enterprises, but this figure fell to 20% in 2010. Reflecting China's demographic trends, the average age of the full sample increased from 35 years in 2001 to 37.5 years in 2010. The average age of non-migrants was generally 5−6 years older than migrants in each of the three years.

As discussed in the "Introduction" section, average years of schooling for the full sample and the migrant and non-migrant subsamples increased from 2001 to 2010. In

Table 2. Descriptive statistics.

	2001			2005			2010		
	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Number of observations	6197	3374	2823	6490	3340	3150	8956	4542	4414
Gender (%)									
Male	57.24	55.50	59.31	56.29	57.29	58.23	55.30	53.52	57.13
Female	42.76	44.50	40.69	43.71	42.71	41.77	44.70	46.48	42.87
Marital status (%)									
Unmarried	24.92	24.92	37.16	16.50	16.54	18.55	19.25	16.39	22.02
Married	75.08	75.08	62.84	83.50	83.46	81.45	80.75	83.61	77.98
Age (years)	35.09	39.61	29.68	41.89	45.29	38.68	37.49	40.00	35.05
Age distribution (%)									
16–24 years	19.78	8.33	33.48	2.25	0.44	3.95	10.40	2.00	11.33
25–34 years	30.51	22.55	40.03	22.79	17.04	28.25	31.52	20.59	45.75
35–44 years	27.84	35.18	19.06	34.04	24.18	43.33	30.29	39.53	26.79
45–54 years	18.11	28.13	6.13	29.61	40.23	19.60	22.55	20.62	9.12
55–64 years	3.76	5.81	1.31	11.31	18.12	4.88	5.24	17.26	7.00
Years of Schooling	10.39	11.89	8.60	10.44	12.01	8.96	11.46	12.61	10.33
Education attainment (%)									
Primary or below	10.69	2.94	21.34	11.48	1.94	20.46	6.73	5.46	15.19
Junior Middle	39.99	25.85	53.46	38.14	23.19	52.22	33.34	27.49	35.42
Senior Middle or Polytechnic	31.05	42.19	21.13	34.62	47.57	22.41	33.06	29.87	30.74
Three Year College	9.97	16.69	2.80	8.90	15.49	2.70	14.80	29.78	15.52
Four Year University or above	8.30	12.33	1.28	6.86	11.80	2.22	12.07	7.40	3.13
Years of Post School Experience	18.13	21.13	14.50	24.45	26.29	22.71	19.30	20.58	18.04
Residence registration (*Hukou*) (%)									
Rural Hukou	63.09	1.81	78.91	56.78	2.22	81.81	59.20	3.78	76.80
Urban Hukou	36.91	98.19	21.09	43.22	97.78	18.19	40.80	96.22	23.20

(*continued*)

Table 2. (*Continued*)

	2001			2005			2010		
	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Employment status (%)									
Employees	71.99	89.51	51.07	60.93	87.95	35.98	71.22	87.20	55.04
Self-employed	28.01	10.49	48.93	39.07	12.05	64.02	28.78	12.80	44.96
Unit type (%)									
Government Agency and Institution	18.98	29.02	6.95	14.32	24.54	4.73	14.21	23.95	4.74
State-Owned/Collective Enterprise	27.71	43.56	8.73	19.49	33.57	6.20	19.82	32.19	7.80
Private Enterprise	11.88	8.18	16.32	14.04	17.13	11.14	23.65	21.19	26.02
Foreign/Taiwan/HK JV	2.87	3.78	1.78	2.33	3.45	1.26	3.29	3.60	3.04
Individual Business	37.76	14.42	65.73	45.20	15.44	73.23	34.72	16.22	52.68
Others	0.79	1.04	0.50	4.62	5.87	3.43	4.31	2.85	5.72
Industry (%)									
Manufacturing	15.63	22.04	7.74	12.40	19.85	5.43	11.16	15.03	7.40
Construction	6.55	5.00	8.46	3.64	2.85	4.38	5.15	3.76	6.54
Transportation/Post/ Communication	7.10	10.53	2.88	7.09	11.10	3.32	11.21	15.07	7.45
Wholesale/Retail/Hotel/Catering	27.05	14.70	42.24	37.13	20.19	52.96	33.17	21.11	44.88
Social Service	19.17	10.81	29.45	22.94	18.73	26.86	19.49	14.52	24.31
Others	24.50	36.92	9.22	16.81	27.28	7.05	19.82	30.51	9.42
City (%)									
Shanghai	21.04	22.26	19.59	23.16	23.22	23.07	18.69	16.64	20.67
Wuhan	20.36	21.49	19.02	19.21	18.40	19.99	23.62	25.11	22.15
Shenyang	18.82	15.03	23.34	17.32	17.86	16.85	18.01	17.82	18.19
Fuzhou	20.36	22.05	18.35	22.08	22.59	21.60	19.00	20.79	17.31
Xi’an	19.41	19.18	19.70	18.23	17.93	18.49	20.68	19.63	21.69

Source: Authors’ calculations using data from the China Urban Labor Survey (2001, 2005, 2010).

Table 3. Educational attainment and hourly wage (in 2010 RMB).

	2001			2005			2010		
	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Urban	Migrant
Primary or below	3.99	4.79	3.85	3.89	4.15	3.87	6.78	8.22	6.53
Junior Middle	4.70	5.57	4.19	4.75	5.53	4.42	7.76	8.19	7.56
Senior Middle or Polytechnic	6.45	6.79	5.67	6.78	7.07	6.21	10.13	10.70	9.30
College	8.92	8.91	8.98	9.77	9.76	9.83	14.00	14.22	13.53
University or above	12.17	11.70	17.57	14.17	14.19	14.09	19.96	19.20	21.79
Total	6.17	7.37	4.73	6.45	7.92	5.07	10.87	12.33	9.45

Note: Real wages are attained by using the fixed-base urban CPI to deflate nominal wages.
Source: Authors' calculations using data from the China Urban Labor Survey (2001, 2005, 2010).

each of the three years, the average year of schooling is higher for non-migrants than for migrants. Table 3 shows the relationship between educational attainment and real wage per hour for the full sample and subsamples. The real wage rate increased by 75% for the full sample, 67% for non-migrants and 100% for migrants between 2001 and 2010. Wages for non-migrants exceed that for migrants in each of the three years. For the full sample as well as each subsample, there is a positive relationship between educational attainment and wages in each of the three years.

Empirical specification and methodology

We employ a Mincer earnings function in which the log of hourly earnings (measured in RMB) is regressed on years of schooling, post-school experience, post-school experience squared and a series of control variables. The specific control variables that we employ are gender, marital status, household registration (*hukou*) and ownership, sector and city dummy variables. A problem with the OLS estimates of the earnings function is that presence of measurement error and omission of an individual's ability may bias estimates of returns to schooling. Thus, in addition to OLS, we also present two-stage least squares (TSLS) estimates in which we instrument for education. The problem with IV estimation is finding a valid IV. The CULS has data on three potential candidates for IVs that have been employed in previous studies to instrument for education; namely spouse's education, father's education and mother's education. Of these three potential IVs, we used spouse's education to instrument for education. Mother's education and father's education were not significant in the first-stage regression and thus were not valid IVs in our case.

Spouse's education has been used as an instrument for education in several previous studies (see e.g. Arabsheibani and Mussurov 2007; Di Pietro and Pedace 2008; Lall and Sakellariou 2010; Trostel, Walker, and Wooley 2002), including studies for urban China (Chen and Hamori 2009; Mishra and Smyth 2013). For spouse's education to be a valid IV, it must be correlated with schooling, but not with the residual in the earnings function. The rationale for using spouse's education as an IV is predicated on the notion of assortative matching, which suggests married couples share common experiences and interests and that many of their characteristics, including years of schooling, will be positively correlated. A number of studies document empirical support for the existence of assortative mating in marriage (see e.g. Pencavel 1998; Weiss 1999).

However, one may not be convinced that spouse's education satisfies the exclusion restriction. The problem is that intuitively it is conceivable that the education of one's spouse will have direct effects on one's wage.[4] Kang and Peng (2012) argue that spouse's education is correlated with family background and therefore not really exogenous. One's family background may influence one's income if it assists one to find a better paying job through connections or nepotism or serve as a proxy for children's productivity, which is not captured by other observables such as education. Chen and Feng (2011) provide evidence consistent with both ideas in urban China.

Given these concerns about the validity of spouse's education as an IV, we used a different identification strategy as a robustness check that does not rely on the assumption that spouse education is uncorrelated with the error term in the wage equation. Specifically, we employ the methodology proposed by Lewbel (2012), which is useful for applications where other sources of identification, such as instrumental variables, are either not available or potentially weak.

The estimation problem of current study can be summarized as

$$Y_1 = X^I ß_1 + Y_2 \Upsilon_1 + \xi_1 \quad \xi_1 = \alpha_1 U + V_1 \tag{1}$$

$$Y_2 = X^I ß_2 + \xi_2 \quad \xi_2 = \alpha_2 U + V_2. \tag{2}$$

Assume Y_1 is wages, Y_2 is schooling and that U denotes the individual's unobserved ability, which affects both his, or her, schooling and productivity. V_1 and V_2 are idiosyncratic errors. The Lewbel (2012) method uses heteroskedasticity in the data to estimate the IV regression (see Baum et al. 2013). Lewbel (2012) suggests that one can take a vector Z of observed exogenous variables and use $[Z - E(Z)]\xi_2$ as an instrument if

$$E(X\xi_1) = 0, \; E(X\xi_2), \quad \text{cov}(Z, \xi_1, \xi_2) = 0 \tag{3}$$

and there is some heteroskedasticity in ξ_j. The intuition behind why $[Z - E(Z)]\xi_2$ works as an instrument is that identification occurs by having regressors that are not correlated with the product of the heteroskedastic errors. The point is that the vector Z could either be a subset of X or equal to X. Using the above-chosen set of instruments, one can use TSLS to estimate the IV regression, as one would do with conventional IVs. Lewbel (2012) also suggests, that in cases where instruments (such as spouse's education) exist, we can estimate Equations (1) and (2) using TSLS with spouse's education and an estimate of $[Z - E(Z)]\xi_2$ as instruments. As ξ_2 is a population parameter, and it cannot be directly observed, we use its sample estimate $\hat{e}_2$, obtained from the first stage regression, and consequently use the vector $[Z - E(Z)]\hat{e}_2$ as IVs.

The Lewbel (2012) approach rests on some important assumptions. It is important to be aware of these (see Baum et al. 2013 and Mishra et al. 2014 for a fuller discussion of these points). The first assumption is that there is heteroskedasticity in ξ_j. The exact form of heteroskedasticity requirement as derived in Lewbel (2012) is cov $(Z, \xi_2^2) \neq 0$. In practice, as an approximation, Lewbel (2012) suggests using the estimate of the sample covariance between Z and squared residuals from the first-stage regression linear regression on X to test for this requirement, using the Breusch and Pagan test for heteroskedasticity. As noted by Lewbel (2012, 71), "… if cov (Z, ξ_2^2) is close to or equal to zero, then $[Z - E(Z)]\xi_2$ will be a weak or useless instrument, and this problem will be evident in the form of imprecise estimates with large standard errors". It is easy enough to test for this assumption. The Breusch and Pagan test for heteroskedasticity is significant and the assumption satisfied.

Second, more generally the assumptions specified in Equation (3) are all based on population parameters and are non-testable. This is not as problematic, however, as it might seem, given the assumptions are all fairly standard. As Lewbel (2012, 69) puts it: "These are all standard assumptions, except that one usually either imposes homoscedasticity or allows for heteroskedasticity, rather than requiring heteroskedasticity". This means, therefore, the only nonstandard required assumption by Lewbel (2012) is heteroskedasticity and this can be tested for in any case.

A further consideration is that the Lewbel (2012) estimates are potentially sensitive to the choice of *Z*. The problem is there are no accepted guidelines or rules of thumb for the particular choice of *Z*. This can leave the analysis potentially open to the critique that one chooses a particular choice of *Z* to get a set of results that best fit the story one is trying to tell. This problem, though, can be addressed through doing a series of robustness checks for different combinations of Z. The main results, reported below, are based on $Z =$ all of *X*, which is common in the literature; however, in robustness checks, which are not reported, we also tested for $Z =$ various subsets of *X*. We find that the results are generally not sensitive to the specific choice of Z and that the heteroskedasticity assumption is met, no matter what subset of *X* is chosen as *Z*.

A final issue is that Lewbel's (2012) estimates are based on higher order moments. As such, it is likely that they are not as reliable as conventional instruments that meet all standard exclusion restrictions. As Lewbel (2012, 67) states: "The resulting identification is based on higher moments and so is likely to produce less reliable estimates than identification based on standard exclusion restrictions, but can be useful in applications where traditional instruments are not available or could be used along with traditional instruments to increase efficiency." Lewbel's (2012) own empirical example, which is an application to estimating the Engel curve, plus studies which have applied the method to estimate returns to schooling and in other contexts suggest that the IV estimates are similar to those using conventional valid IVs.[5]

Results

OLS estimates

Table 4 presents the OLS estimates of returns to schooling for the full sample as well as for migrants and non-migrants separately for each of the three years. For the full sample the returns to schooling increased from 6.8% in 2001 to 7.8% in 2005 and 8.6% in 2010. These estimates are lower than the OLS estimates for urban China reported in some studies for the 2000s (see e.g. Ge and Yang 2011; Qian and Smyth 2008; Zhang et al. 2005). However, they are similar to the 7%−8% range reported in several other recent studies for the years 2004 to 2007 (see e.g. Chen and Hamori 2009; Mishra and Smyth 2014; Ren and Miller 2012). The OLS estimates here are also consistent with predictions for the 2000s based on the meta-analysis results for urban China reported in Liu and Zhang (2012). Moreover, they are much higher than the estimates reported for the period prior to the mid-1990s.

In each of the three years, the returns to schooling for non-migrants were roughly double that of migrants. This finding is consistent with previous studies that have found the returns to schooling are higher for non-migrants than migrants in urban China (see e.g.

Table 4. Results of OLS regressions.

	2001			2005			2010		
Variables	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Years of schooling	0.0678***	0.0869***	0.0461***	0.0781***	0.1138***	0.0467***	0.0860***	0.1183***	0.0632***
	(21.9133)	(22.6074)	(8.6505)	(23.6297)	(22.9276)	(10.3967)	(30.6463)	(28.2321)	(16.0824)
Experience	0.0002	0.0038	0.0060	0.0001	0.0096*	0.0067	0.0065***	0.0139***	0.0017
	(0.0744)	(0.9967)	(1.0495)	(0.0382)	(1.7382)	(1.3064)	(2.7084)	(4.0550)	(0.5068)
Experience squared	0.0000	−0.0001*	−0.0000	−0.0000	−0.0001	−0.0002*	−0.0002***	−0.0004***	−0.0002**
	(0.4632)	(−1.6614)	(−0.2562)	(−0.5956)	(−1.4043)	(−1.8810)	(−4.8234)	(−4.7932)	(−2.2784)
Male	0.1624***	0.1606***	0.1592***	0.1716***	0.2155***	0.1458***	0.1822***	0.1735***	0.2309***
	(9.8282)	(8.4956)	(5.5239)	(11.3519)	(10.1022)	(7.0085)	(14.8566)	(10.1025)	(12.0653)
Married	0.1108***	0.0243	0.1474***	0.0427	0.1411***	0.0078	0.1804***	0.1459***	0.2361***
	(3.9311)	(0.6665)	(3.2570)	(1.4842)	(3.1767)	(0.2133)	(8.7398)	(4.8959)	(7.4203)
Rural registration	−0.2693***	−0.1577**	−0.3373***	−0.1213***	−0.0829	−0.2875***	−0.0591***	−0.0922**	−0.2976***
	(−12.3231)	(−2.3346)	(−9.6859)	(−6.1441)	(−1.1869)	(−10.3996)	(−3.8391)	(−2.1006)	(−11.8274)
Constant	0.9567***	0.9102***	0.9466***	1.3432***	1.1728***	1.3927***	1.6454***	1.3658***	1.7682***
	(15.4623)	(12.4770)	(8.5072)	(19.3645)	(12.0568)	(13.6469)	(31.1594)	(18.1958)	(22.6567)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Unit-type dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
City dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	5685	3162	2523	6196	2992	3204	8706	4309	4397
R^2	0.3714	0.3396	0.2526	0.3990	0.4609	0.2792	0.4207	0.4146	0.3946

Notes: t-statistics in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Maurer-Fazio and Dinh 2004). Returns to schooling increased over the decade for both migrants and non-migrants. For non-migrants there was a jump between 2001 and 2005, but returns to schooling remained fairly stable between 2005 and 2010. This result is consistent with Liu and Zhang's (2012) conclusion that the returns to schooling for urban workers slowed in the Global Financial Crisis. For migrants, returns to schooling were constant between 2001 and 2005, but increased between 2005 and 2010. The latter likely reflects periods of shortage in migrant skilled labor, and resultant wage increases, in urban coastal areas since 2004.

For most of the models, the coefficient on experience and experience squared is insignificant, although the earnings−experience profile is significant with the expected parabolic shape for the full sample and the non-migrant sample in 2010. We find that males earn more than females. For the full sample, the gender wage gap shows a slight increase from 16.2% to 18.2% between 2001 and 2010. The gender wage gap for migrants and non-migrants was similar in 2001 (15.9%−16%), but differed in 2005 and 2010. In 2005 it was larger for non-migrants than migrants and in 2010 this was reversed. Marital status is significant in two-thirds of the regressions. For the full sample the premium on being married increased from 11.1% in 2001 to 18% in 2010. This result is consistent with increasing cross-productivity effects within marriage over time (Birch and Miller 2006).

The coefficient on rural household registration is negative and significant in all regressions, although for the full sample it has decreased from 26.9% in 2001 to 5.9% in 2010. Those with a rural household registration constituted 75%−80% of the migrant sample in each of the three years. These individuals earned roughly one third less than their counterparts who had acquired an urban household registration over the course of the decade. A small percentage of non-migrant households had a rural household registration. This group constituted 1.8% of non-migrants in 2001 and increased to 3.8% of non-migrants in 2010.[6] These individuals incurred a wage penalty of 15.8% in 2001, which declined to 9.2% in 2010, relative to non-migrants who had an urban household registration.

TSLS estimates

The results for the first stage regression and TSLS regression, in which spouse's education is used to instrument for education, are reported in Tables 5 and 6, respectively. Roughly 75%−80% of the sample is married over the three years, so we lose relatively few observations using spouse's education as an IV. The coefficient on spouse's education is positive and significant in the first stage results. The first stage Stock and Yogo (2002) *F*-test, reported at the bottom of Table 6, confirms that spouse's education is a valid IV. The endogeneity test, also reported at the bottom of Table 6, rejects the null hypothesis that the OLS estimates are consistent.

For the full sample, the TSLS estimates for returns to schooling are 8.2% in 2001, 8.7% in 2005 and 9.1% in 2010. For the non-migrant sample, the TSLS estimates for returns to schooling are 11.4% in 2001, 12.4% in 2005 and 11.7% in 2010. For the migrant sample, the TSLS estimates of the returns to schooling are 6.9% in 2001, 5.4% in 2005 and 7.1% in 2010. Generally, the TSLS estimates of the returns to schooling are higher than the OLS estimates, suggesting the OLS estimates exhibit downward bias, but the difference is not large. The largest difference is 2.7% for non-migrants in 2001.

Overall, the TSLS estimates for returns to schooling are slightly lower than most previous studies which have used spouse's education to instrument for schooling. Using spouse's education as an IV, Chen and Hamori (2009) found that the TSLS estimates of returns to schooling in urban China were 12.6% for married males and 14.5% for married

Table 5. First-stage regression results.

	2001			2005			2010		
Variables	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Spouse's education	0.6395***	0.4713***	0.3003***	0.6204***	0.5243***	0.7259***	0.7324***	0.6910***	0.7520***
	(53.0330)	(28.5379)	(12.5540)	(67.7276)	(36.6717)	(62.4297)	(98.6300)	(56.9622)	(79.7505)
Experience	−0.0646***	−0.0017	−0.1427***	−0.0872***	−0.1105***	−0.0528***	−0.0344***	−0.0300**	−0.0387***
	(−4.2658)	(−0.2976)	(−4.6504)	(−6.9282)	(−5.1864)	(−3.5992)	(−4.8273)	(−2.5267)	(−4.2306)
Experience squared	−0.0001	−0.0001	0.0002	0.0002	0.0003	0.0001	−0.0002*	−0.0006**	−0.0000
	(−0.2433)	(−0.5808)	(0.3150)	(0.8549)	(0.7725)	(0.2052)	(−1.6707)	(−2.3990)	(−0.0521)
Male	0.9121***	0.0576**	1.5668***	0.3839***	0.2920***	0.4948***	0.5786***	0.5864***	0.5954***
	(13.9124)	(2.4125)	(11.1435)	(7.4541)	(3.5292)	(8.0897)	(16.1947)	(10.4039)	(12.8372)
Rural registration	−1.1112***	−0.0250	−1.3240***	−1.2341***	−0.7462***	−0.4870***	−0.6138***	−0.5040***	−0.4908***
	(−12.0636)	(−0.2844)	(−7.4444)	(−19.9985)	(−2.7094)	(−6.2483)	(−13.6194)	(−3.5135)	(−7.7996)
Constant	4.8812***	0.4007***	10.0497***	7.1564***	9.5002***	4.1992***	3.5656***	4.5665***	2.8642***
	(11.0705)	(2.8353)	(3.8530)	(34.3637)	(31.4710)	(14.5900)	(21.7960)	(14.8683)	(13.2631)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Unit-type dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
City dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3748	2219	1529	4488	1980	2508	6717	2819	3898
R^2	0.7261	0.4699	0.4353	0.7564	0.6272	0.7814	0.8124	0.7598	0.8107

Notes: t-statistics in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

Table 6. Result of two-stage least squares regression using spouse's education as an instrument.

	2001			2005			2010		
Variables	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Years of schooling	0.0824^{***}	0.1139^{***}	0.0693^{***}	0.0872^{***}	0.1243^{***}	0.0540^{***}	0.0905^{***}	0.1170^{***}	0.0711^{***}
	(14.5611)	(5.2546)	(15.2449)	(17.6755)	(14.9826)	(8.6275)	(25.0430)	(19.6463)	(15.3930)
Experience	0.0120^{***}	0.0127^{**}	0.0051	0.0019	0.0139^{*}	0.0062	0.0003	0.0044	0.0059^{*}
	(2.6983)	(2.5463)	(0.5499)	(0.3936)	(1.8551)	(1.0364)	(0.1223)	(0.9721)	(1.6626)
Experience squared	-0.0004^{***}	-0.0003^{***}	-0.0004^{**}	−0.0000	-0.0003^{**}	-0.0002^{*}	−0.0001	−0.0001	−0.0000
	(−3.8913)	(−3.0591)	(−2.4804)	(−0.3540)	(−2.1620)	(−1.6875)	(−1.5104)	(−1.5107)	(−0.0629)
Male	0.1354^{***}	0.1775^{***}	0.0317	0.1921^{***}	0.2528^{***}	0.1596^{***}	0.2320^{***}	0.2342^{***}	0.2346^{***}
	(6.4773)	(7.9548)	(0.6778)	(10.1238)	(8.9855)	(6.3565)	(16.6391)	(10.7167)	(12.9654)
Rural registration	-0.1742^{***}	-0.1562^{*}	-0.1619^{**}	-0.2083^{***}	-0.1743^{*}	-0.3004^{***}	-0.1638^{***}	-0.1511^{***}	-0.1744^{***}
	(−5.5259)	(−1.9099)	(−2.5129)	(−8.2808)	(−1.8297)	(−9.2543)	(−9.0293)	(−2.6901)	(−6.8892)
Constant	1.0528^{***}	1.2382^{***}	0.0605	0.8828^{***}	0.5406^{***}	1.2019^{***}	1.3604^{***}	0.9656^{***}	1.6335^{***}
	(7.0366)	(9.5563)	(0.0756)	(10.1541)	(3.7132)	(11.3593)	(21.7919)	(8.3053)	(21.2803)
Industry dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Unit-type dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
City dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	3921	2329	1592	4786	2129	2657	7212	3142	4070
R^2	0.3261	0.3241	0.1832	0.3198	0.3676	0.2110	0.3740	0.3416	0.3562
Diagnostic tests for IV estimation									
Exogeneity χ^2 test	35.14^{***}	6.31^{**}	12.31^{***}	21.34^{***}	8.97^{***}	8.12^{***}	13.54^{***}	4.76^{***}	10.34^{***}
First stage (Stock & Yogo) F-test	308.60^{***}	449.16^{***}	168.77^{***}	479.33^{***}	148.97^{***}	385.49^{***}	1040.51^{***}	357.16^{***}	667.77^{***}

Notes: t-statistics in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.1$.

females in 2004−2006. Similarly, using spouse's education as an IV, Mishra and Smyth (2013) found TSLS estimates of the returns to schooling in the range 21%−23% for 2007. However, it is important to note that Chen and Hamori's (2009) sample was restricted to urban non-migrant households so their TSLS estimates are only slightly larger than our estimates for non-migrant households for 2005. Mishra and Smyth (2013) focused on ethnic Koreans, living in China's northeast, who may not be representative of the urban population as a whole.

While the diagnostics suggest that spouse's education is a valid instrument, one may still be worried that spouse's education is correlated with one's earnings and the exclusion restriction is fundamentally not testable. Thus, for the purposes of robustness, Table 7 reports TSLS estimates using the Lewbel (2012) method. The Breusch−Pagan test for heteroskedasticity rejected the null of constant variance in each case, which is a precondition to apply the Lewbel (2012) method. The top panel reports TSLS estimates for returns to schooling using $[Z - E(Z)]\xi_2$ alone as an instrument. The second panel uses $[Z - E(Z)]\xi_2$ and spouse's education as instruments. The diagnostics tests suggest that instrument validity is satisfied. For both sets of results, the first stage F-test is satisfied and the endogeneity test rejects the null hypothesis that the OLS estimates are consistent. When using both $[Z - E(Z)]\xi_2$ and spouse's education as IVs, because the model is over-identified, we calculate the Sargan chi-square statistic to test for instrument exogeneity. The results, reported at the bottom of Table 7, suggest that the instruments are exogenous.

The TSLS estimates for returns to schooling using $[Z - E(Z)]\xi_2$ alone as an instrument are generally slightly lower than the OLS estimates. As expected, generally the TSLS estimates for returns to schooling using $[Z - E(Z)]\xi_2$ and spouse's education as instruments lie between the TSLS estimates using each IV on its own and, in most cases, are very close to the OLS estimates. Our results are consistent with Sabia (2007) who found that TSLS estimates using the Lewbel (2012) methodology were much lower than TSLS estimates using a conventional IV and were similar in magnitude to the OLS estimates.[7] Overall, for the full sample, the OLS results and the TSLS results using alternative identification strategies from lower to upper bound estimates are within 3 percentage points of each other for 2001 and 2005 and within 2 percentage points of each other for 2010. Each estimate suggests a slight upward trajectory in returns to schooling over the course of the decade.[8]

The main limitation of our Lewbel (2012) results is that the Lewbel (2012) method relies on a set of assumptions that stem from unobservable population parameters that are not testable. Having said this, the data is characterized by heteroskedasticity, which is a precondition to implement the Lewbel (2012) test, and our results are not sensitive to the particular choice of Z. A further limitation is that because the Lewbel (2012) method relies on higher order moments the IV estimates may not be as reliable as those derived with strong conventional IVs. Here, the TSLS estimates using $[Z - E(Z)]\xi_2$ and spouse's education as instruments plausibly lie between the OLS and conventional TSLS estimates. Thus, the Lewbel (2012) estimates serve their purpose as a robustness check on our conventional IV and provide added confidence in the results using spouse's education as a conventional IV.

The findings in Table 8 extend the earlier results to consider returns to schooling for males and females separately. The average years of schooling of both males and females increased in the first decade of the twenty-first century. The average years of schooling of males increased from 10.45 years in 2001 to 11.54 years in 2010, while the corresponding increase for females was 10.30 years to 11.35 years.

Table 7. Estimates of the coefficient on years of schooling using $(z-\bar{z})\hat{\varepsilon}_2$ and spouse's education as instruments.

	2001			2005			2010		
Variables	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant	Full	Non-migrant	Migrant
Two-stage least squares using $(z-\bar{z})\hat{\varepsilon}_2$ as an instrument									
Years of schooling	0.0519**	0.0802***	0.0440***	0.0560**	0.0916***	0.0478***	0.0721***	0.0955***	0.0514***
	(2.3815)	(19.6920)	(6.8480)	(1.9758)	(17.6371)	(4.8283)	(5.1861)	(21.8712)	(10.5003)
Diagnostic tests for IV estimation									
Exogeneity χ^2 test	34.75***	3.07*	10.50***	4.22**	3.12*	10.96***	6.71***	4.32**	3.03*
First-Stage F-test	135.27***	216.47***	214.80***	347.83***	336.78***	219.76***	344.13***	323.58***	243.76***
Two-stage least squares using $(z-\bar{z})\hat{\varepsilon}_2$ and spouse's education as instruments									
Years of schooling	0.0612***	0.0738***	0.0442***	0.0725***	0.0976***	0.0430***	0.0849***	0.0969***	0.0608***
	(15.6479)	(16.4382)	(6.8075)	(14.7260)	(16.1540)	(5.9446)	(23.8326)	(19.1843)	(14.2091)
Diagnostic tests for IV estimation									
Exogeneity χ^2 test	3.08*	3.14**	2.901*	12.78***	3.12**	4.33***	3.97**	2.97*	3.17**
First-Stage F-test	259.36***	320.47***	395.48***	253.29***	678.49***	347.83***	546.21***	191.62***	544.13***
Sargan χ^2 test	1.058	0.838	0. 698	1.253	0.767	0.694	1.289	0.581	0.771

Notes: t-statistics in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$. Each specification is estimated with full set of controls, including city, industry and unit-type dummies.

Table 8. Estimates of the coefficient on years of schooling for full sample for males and females.

	OLS		TSLS (Spouse IV)		TSLS ($(z-\bar{z})\hat{\varepsilon}_2$ IV)		TSLS (Spouse & $(z-\bar{z})\hat{\varepsilon}_2$ IVs)	
	Male	Female	Male	Female	Male	Female	Male	Female
2001	0.0703***	0.0626***	0.0858***	0.0788***	0.0557***	0.0464***	0.0631***	0.0586***
	(18.3101)	(15.8904)	(9.6958)	(11.9963)	(14.2225)	(10.7578)	(13.0957)	(10.1297)
2005	0.0816***	0.0692***	0.0940***	0.0792***	0.0631***	0.0528***	0.0803***	0.0615***
	(19.3362)	(15.1454)	(14.0456)	(10.4743)	(16.0471)	(11.2687)	(13.397)	(8.3862)
2010	0.0927***	0.0789***	0.1030***	0.0850***	0.0826***	0.0659***	0.0927***	0.0735***
	(24.5874)	(21.2634)	(18.8638)	(16.226)	(20.1248)	(16.8567)	(18.2332)	(14.9171)

Notes: t-statistics in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$. Each specification is estimated with full set of controls, including city, industry and unit-type dummies.

We report the returns to schooling for males and females for the full sample for each of the three waves of the CULS, using OLS and TSLS with each of the alternative identification strategies used earlier. Two features of the results are that returns to schooling increased for both males and females over the course of the decade and that the returns to schooling were higher for males than for females. The latter findings are inconsistent with the conventional wisdom on gender differentials in returns to schooling in urban China. However, they are consistent with recent results reported in Chen and Hamori (2009), Qian and Smyth (2008), Kang and Peng (2012) and Ren and Miller (2012) who use a range of datasets for urban China.[9]

Discussion

The main finding is that the premium to education increased over a period in which education levels were rapidly expanding. This result suggests that the strong demand for skilled labor and workplace innovations increased the returns to education, despite the substantial increase in the number of post-secondary graduates. China has traditionally been considered to have an abundance of low-cost labor. China is now aiming to move up the value-added ladder and become a player in more complex technologies. This generates increased demand for skilled labor, which is increasing the returns to schooling. The returns to skilled labor increased in the range of 15%−20% per annum in the first decade of the twenty-first century (Boston Consulting Group 2011). There is evidence that wages in IT in the big cities on the Chinese coastal seaboard are approaching those available in Europe and the United States. Some of the big Chinese firms such as Datang, Huawei and Lenovo are paying salaries in the same range as those available in the United States, in a bid to attract skilled workers to their engineering design units and R&D laboratories (Simon and Cao 2008).

The overall results for returns to schooling over the first decade of the twenty-first century are consistent with rising urban wage inequality between skilled and unskilled labor over the same period. Gustafsson, Li, and Sicular (2008, 1) state: “Income inequality is . . . . now considered high by international standard. [I]n China, the speed with which the increase has occurred, and the level to which inequality has arisen, is striking.” Overall wage inequality in China increased between 2000 and 2010 based on different measures. For example, the Gini coefficient increased from 0.34 to 0.37 and the ratio of the 90th to the 10th percentiles increased from 4.96 to 5.59 (Meng, Shen, and Sen 2013).

Turning to the differences between migrants and non-migrants, while both groups experienced an increase in average years of schooling, the average years of schooling is higher for non-migrants than for migrants. At the same time, the increase in the returns to schooling is higher for non-migrants than for migrants over the decade. There are several possible explanations for the relatively low returns to migrants, centering on segmentation in the urban labor market between migrants and non-migrants.

One possible explanation for this finding is differences in educational quality between urban schools on the one hand and migrant and rural schools on the other. A second potential explanation is that it represents discrimination against migrants (Meng and Zhang 2001). A third possibility is that migrant workers might need to accumulate experience in urban areas before they can enjoy the same returns as non-migrant workers. A fourth possibility is that most migrants take jobs in low-skilled occupations in construction and manufacturing. Highly educated workers employed in low-skilled jobs may experience low returns to education (Cui, Nahm, and Tani 2013).

Disparity in average years of schooling and returns to schooling are also reflected in within-group differences in income inequality between non-migrants and between migrants. Zhao and Qu (2013) find that over the period of 2002−2007, non-migrants experienced substantial growth in wage inequality in urban China, while wage inequality among migrants decreased over the same period. The reason wage inequality decreased among migrants was that high-wage migrants experienced slower wage growth than low-wage migrants. Zhao and Qu (2013) interpret this result as migrant workers experiencing a glass ceiling in the urban labor market.

In the results for males and females, we find that returns to schooling for both genders increased over the course of the decade and that the returns to schooling were higher for males than for females. The latter can be explained in terms of female educational expansion and rising participation rates. As suggested by Ren and Miller (2012), the rapid growth in the number of female graduates since 2000 has served to negate the undersupply of highly skilled female professionals in urban China.

In 2010, women represented half of undergraduate students and just under half of postgraduate students enrolled in Chinese universities (OECD, 2011). At the same time, women make up just under four in every 10 MBA students at the top-ranked Chinese programs; figures that are similar to those at the leading business schools in the United States (Hewlett, 2011). This has resulted in an over-supply of highly skilled females, relative to highly skilled male graduates in China. Li, Whalley, and Xing (2014) find that following the expansion in higher education in China, a higher proportion of female than male graduates were unable to find positions.

Conclusion

The purpose of this paper was to examine whether the expansion in higher education, and associated increase in average years of schooling, that has occurred in urban China since 2000 has reduced the premium to education. To address this issue, we employed three waves of the CULS for 2001, 2005 and 2010. The main finding was that the estimated premium to education increased hand-in-hand with the expansion in the higher education system in China. This finding is consistent with the demand for skilled labor outstripping the increase in supply as China moves into higher value-added production. Drilling down from the overall results, to focus on the results for migrant and non-migrant samples, we find that the average years of schooling and returns to schooling for non-migrants are higher than migrants. This result is consistent with continuing segmentation in the Chinese urban labor market.

In terms of the public policy implications of our results, first our findings suggest that simply expanding education access will not improve social equity. Our finding that education expansion is associated with an increase in the returns to schooling is consistent with increased wage inequality in the urban labor market. Possible policies to address this issue include reduction in tuition fees and measures to reduce the financial burden on students, which increased through the 2000s (Heckman, 2005). A reduction in tuition fees would reduce wage inequality by increasing access to education for low-income individuals and increasing their returns to education in the form of higher wages in the future compared with the counterfactual.

Second, the Chinese government has long emphasized the importance of gender equality. Our results suggest that the gender wage gap increased throughout the 2000s. Returns to schooling are an important factor of wage determination. That we find returns to schooling were lower for females than males could help explain the observed increase in gender inequality. Third, that returns to schooling for migrants are lower than their non-migrant counterparts ought to be worrying for policy makers at a time when this particular group represents 44% of the urban labor market. It suggests that measures to improve the quality of schools migrants attend as well as the human capital base of migrants are needed to increase their returns to education.

We conclude by referring to some of the limitations of our results and suggestions for future research. First, the CULS provides a snapshot at the beginning of the decade, middle of the decade and end of the decade. Future research could examine returns to schooling in the intervening years. Second, the results are restricted to five of China's largest cities. Future research could examine the returns to schooling in urban China since 2000 using a broader set of cities. Third, the CULS picks up more educated migrants who are successful at finding employment in larger cities (de Brauw and Giles 2006). This suggests that the returns to education reported here represent an upper bound on the returns to schooling for migrants in China's urban labor markets. This is an issue that could be fruitfully explored using alternative datasets.

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Notes

1. We use the terms 'migrant' and 'non-migrant' as opposed to 'migrant' and 'urban' because some migrants in the survey have acquired an urban household registration and some non-migrants, living on the urban fringes, have a rural household registration.
2. The urban household surveys have contained data on rural–urban migrants since 2002, but the data are not representative of this group – see the discussion in Meng, Shen, and Sen (2013).
3. However, in robustness checks, which are reported in the working paper version (Gao and Smyth 2012), we exclude the self-employed and the returns to schooling that are similar to those reported below.
4. For an overview of the literature on the relationship between spouse's education and male earnings, see Birch and Miller (2006).
5. For previous applications of the Lewbel method to estimate the returns to schooling in China, see Mishra and Smyth (2012, 2013).

6. The cities surveyed in the CULS have agricultural districts, which tend to be located on the outskirts of the city proper, in which non-migrant residents have a rural household registration.
7. Sabia (2007) relies on the working paper version of Lewbel (2012).
8. As a robustness check we estimated the coefficients on schooling using different choices of Z, where each choice of Z constituted a subset of *X*. The estimates using different choices of Z are similar in magnitude to using $Z = X$ and were within a couple of percentage points of the equivalent results reported in Table 7. In robustness checks reported in Gao and Smyth (2012) we also estimate the returns to schooling for the full sample using different combinations of industry and unit-type dummies, including omitting all dummies. The estimates for returns to schooling are sometimes slightly higher and sometimes slightly lower than the equivalent results given in Tables 4, 6 and 7 and are generally similar in magnitude. The trend of a slight upward trajectory in returns to schooling remains robust omitting industry and unit-type dummies, reinforcing the results reported here.
9. Earlier studies using the CULS 2005 have also found that returns to schooling are higher for males than for females (see e.g. Gao and Smyth 2010, 2011).

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