CORE
Metadata, citation and similar papers at core.ac.uk
Provided by Munich Personal RePEc Archive

MPRA
Munich Personal RePEc Archive

Siblings, public facilities and education returns in China

Lili Kang and Fei Peng

May 2012

Online at https://mpra.ub.uni-muenchen.de/38922/
MPRA Paper No. 38922, posted 24 May 2012 14:30 UTC

Siblings, Public Facilities and Education Returns in China

Lili Kang†*

Fei Peng*

† Corresponding author. Institute of Economics, School of Humanity and Social Sciences, Tsinghua University, Tsinghua Yuan No. 1, Ming-Zhai, Beijing, 100084, P.R.China. E-mail address: ll.kang@yahoo.com

* Centre for Research on the Economy and the Workplace (CREW), Birmingham Business School, University House, Edgbaston Park Road, Birmingham, B15 2TY, UK.

For useful comments we thank John Knight, Mary O'Mahony, Stan Siebert, Georgios Efthyvoulou and other participants of the CEA conference (2010) at the University of Oxford and the workshop at the University of Birmingham (2010). The China Health and Nutrition Survey (CHNS) data are used with the permission of the Carolina Population Centre based at the University of North Carolina at Chapel Hill. Neither the original collectors of the data nor distributors bear any responsibility for the analyses or interpretations presented here. All remaining errors are our own.

ABSTRACT

This paper investigates the intrahousehold resource allocation on children's education and its earnings consequence in Chinese labour market. In order to overcome the endogeneity problem of schooling, we consider the siblings structure and the available public facilities as instrumental variables. Females' education is negatively affected by siblings (brothers or sisters) number, while males' education is also negatively affected by their brothers but much less by their sisters. For the youngest cohort born after 1980, the education of a girl would be heavily impeded by her sisters, reflecting strong distortion of "One-Child Policy" on intrahousehold resource allocation. Comparing the OLS and GIV estimations for returns to schooling, we find that there are downwards biases of OLS estimations for males in all cohorts and in all years. However, for females, downwards biases of OLS estimation are only for data before 2004, as females in the old cohorts actually have upwards biases after 2004. Education returns of the youngest cohort are much higher than old cohorts supporting the argument of heterogeneous human capital accumulation during transition.

KEYWORDS: Siblings, education returns, China
JEL Codes: J13, J16, J31

I. Introduction

A number of studies find that returns to education increased from the ‘‘pre-transition’’ period to the ‘‘early transition’’ period in transition countries. Rutkowski (2001) argues that the increase in the premium to university education has been the main observable factor behind the increase in earnings inequality during the transition of central European countries. However, Flabbi et al. (2008) find that the rising trend in returns to education over the transition period is actually quite weak in eight transition economies (Bulgaria, Czech Republic, Hungary, Latvia, Poland, Russia, Slovak Republic and Slovenia), even though there are significant differences in returns across transition countries. Thus, the features of the changes of education returns over the transition period still remain vague.

As the largest transition economy in the world, policy makers, academics and the public in China have shown long interest in the changing nature of education returns. As Zhang et al. (2005) point out, education returns provide information about the incentives for human capital accumulation, the efficiency of resource allocation, and the distributional consequences of differences in human capital. Chinese labour market liberalization, more flexible wage setting and the broader transition to a market economy since the 1980s are assumed to lead to an increase in education returns which has far-reaching social, political, and economic gender-differentiated implication (Berik et al., 2007). However, some factors may offset the expected increases of education returns. For example, skills and experience acquired under central planning may be less useful in the new market environment. The pace of privatization, enterprises restructuring and changes in labour force participation of married women may also either exacerbate or offset gender-differentiated rising education returns (Kang and Peng, 2013). In a society where the state has a historic commitment to promoting gender equality (Gal and Kligman, 2000), a study of education returns in China based on a wide view of gender, family and society is highly demanded.

Before the economic reforms in 1978, China had virtually no labour market, and nearly all firms were state-owned enterprises or de facto state-owned enterprises. Wages of all workers were determined and controlled through a rigid system to reduce labour costs for rapid industrialization. Low wages were made possible by state provision of public facilities and services to workers and their families (Peng, 1992; Nee, 1996). This heavy planning led to poor effort incentives which depressed productivity and innovation and made the family income much dependent on public facilities such as energy, water and transportation. There were several wage reform waves in the late 1980s and 1990s to create a more market-oriented wage structure. The former rigid wage system has been replaced by two components consisting of fixed and variable wages.[1] In 1986, the State Council formally introduced individual labour contracts to end the system of permanent employment. Thus, by the early 2000s, the labour market began to decide wages and employment in China (Meng and

[1] The fixed portion includes the basic wages, seniority wage, insurance (medical, unemployment and pensions) and a housing fund. The variable portion includes bonuses, based on both individual productivity and enterprise profitability.

Kidd, 1997; Freeman, 2007).

On the other hand, Chinese education institutions have been transitioning from a highly regulated and rigid system to a market-based and flexible system. In 1986, China announced a compulsory education policy in 1986 and gradually enforced this policy to both urban and rural areas of China in later years. At the same time, China reformed the financial system of higher education,[2] stating that excellent students can apply for scholarships, and students with financial difficulties could receive necessary grants which includes living stipends, government-subsidized student loans, work-study scholarships, grants and tuition waivers (Li, 2007). From 1999, the state launched a large-scale expansion of the higher education sector. In 1998 new enrolments in higher education were 1.08 million, but by 2006, this number had increased to 4.75 million (Qian and Smyth, 2011). Thereafter, the intrahousehold decisions on education investment have been affected by these institutional changes, which provide a rare chance of natural experiment to assess the effects of labour market and education policy on the gender-differentiated education returns in China, concerning the bargaining and competition intrahousehold.

In this paper, we investigate the returns to schooling in China over the period of 1989-2009 using the China Health and Nutrition Survey (CHNS) dataset. We address

[2] On May 27, 1985, the policy document entitled "Decision of the CPC Central Committee Concerning the Reform of the Educational System" began to allow higher education institutions to "enroll a limited number of self-financed students beyond the national enrolment plan".

the question: as China has transformed from a central planned economy to market allocation of resources, do the private returns to education as wage differentials reflect the gender-differentiated productivity associated with education or gender-differentiated education chances based on intrahousehold resources allocation? In China, there is a long-standing social norm that a son is generally preferred than a daughter (Lee and Wang, 1999). Only sons could carry the family name and inherit the family patrimony (Bernhardt, 1995). The "One-Child Policy" in China enforced after 1980 forbid a couple to have more than one child while there are some opt-out articles for ethnical minority and people in the rural area (Cameron et al. 2011), making the gender of these children more important (Edlund et al., 2007; Qian, 2008). As the cohort born after 1980 gradually becomes an important part, and then will be the majority of the largest labour force in the world, understanding on the effects of policies of labour market, education and population on their lives has become an urgent job for labour scholarship in China. Study on education return in this paper would shed some new light on this issue.

This paper adopts the ordinary least square (OLS) and generalized instrumental variable (GIV) methodology concerning the intrahousehold decision on education investment. In order to overcome the endogeneity problem of schooling, we consider the siblings structure and the public facilities as instrumental variables, and assess whether the OLS estimates (assuming schooling an exogenous variable) significantly differ from the GIV estimates that consider this bias. Therefore, this paper contributes to the published literature by investigating the gender-differentiated education returns over the early and later transition periods in China and speculating on the likely institutional and structural factors underpinning these results. The study is the first empirical analysis of its kind for China and provides valuable new insight into the gender-differentiated education returns concerning the policy effects on intrahousehold education decision. The rest of this paper is organized as follows. In section II, we present the OLS and GIV empirical specifications and potential instruments for endogeneity of schooling. Data descriptive statistics and estimates of the returns to schooling are presented in Section III and IV. Section V concludes.

II. Empirical Methodology

Many researchers apply the OLS methodology to examine the rates of returns to education in China. Following this tradition, we first estimate the returns to education based on the Mincerian earning equation commonly applied in the literature (Harmon et al., 2001; Card, 1999; Harmon et al., 2004) as follows:

$$(1) \qquad \ln W_i = \beta_0 + \beta_1 E_i + \beta_2 X_i + \varepsilon_i$$

where lnW_i is the log form hourly wage rate of individual i, E_i represent education attainment, X_i is a vector of control variables including gender, experience and its square, location (urban and province dummies) ; ε_i is an error term with $\varepsilon \sim N(0, \sigma^2)$.

Many authors have addressed the endogenous nature of schooling may bias OLS estimates of returns to schooling (Dearden, 1999; Card, 1999; Blundell et al., 2005). In our CHNS sample, endogeneity can arise because of omitted information of workers' inner ability. That is, the coefficient of returns to schooling β_1 is upwards biased because the chosen schooling levels are positively correlated with omitted ability variable, while ability is positively correlated with the wage rate (Chen and Hamori, 2009). However, Card (1999) argues that OLS estimates of β_1 are biased downwards because individuals with high discount rates choose low levels of schooling, that is, schooling with higher marginal rates of return. In addition, Li and Luo (2004) argue that OLS estimators will be biased downwards caused by measurement error in schooling variable. They estimate returns to education for young workers in urban China and find the OLS estimator are biased downwards by 7-8 percent per year of schooling.

Heckman and Li (2004) recommend using the Instrumental Variable (IV) methodology to cope with the endogeneity problem in China. They use parental education and year of birth as instrumental variables to identify the returns to higher education for young people in the urban areas of the six provinces in 2000. The IV estimator of average return to four-year college attendance is 43% (on average, 11% annually). Moreover, Fleisher and Wang (2005) use birth year, location dummies of elementary education (rural/small town, medium city, or large city), and their interaction as IVs of endogenous schooling and find IV estimated education returns for an additional year of schooling is 10 percent in every year in 1984, 1987 and 1990. Chen and Hamori (2009) use spouse education as instrument and examine education returns in urban China using the pooled CHNS 2004 and 2006 data. Even with these trials, validity of instruments for schooling is still a big problem in studies of education returns and not satisfactory at either theoretical or statistic sense. To be a valid instrumental variable for the endogenous years of schooling, the potential exogenous instruments should be only theoretically and statistically correlated with years of schooling, and not correlated with the unobserved errors in the wage equation $\varepsilon_{i,}$ such as the unobservable inner ability. Belley and Lochner (2007) argue that family background (such as parental income) is correlated with children's cognitive skill and hence is not a valid instrumental variable for children's education. Thus, Heckman and Li (2004) also agree their instrumental variable (parental income) should be put into the wage equation directly as the proxy of ability rather than used as an instrument.

Moreover, parental income is correlated with parental education, so parental education would not be a valid instrument too. Following the same line of reasoning, although birth year used in Fleisher and Wang (2005) may reflect the cohort effect on education, it is possible that the cohort effect is correlated with wage through different channel such as experience variables in the Mincerian equation. Economists have long recognized that people tend to mate with those who have similar individual endowments, such as education and other labor market characteristics (Becker and Lewis, 1973), and China is no exception. Hence, spouse education in Chen and Hamori (2009) is also highly correlated with his/her family background which is generally similar to each other and not really exogenous.

Therefore, this paper tries three potential instrumental variables: number of brothers, number of sisters and access to tap water for education return studies in

China. First, there is an extensive theoretical literature that postulates a trade-off between child quantity and quality within a family, introduced by Becker (1960) and expanded in Becker and Lewis (1973) and Becker and Tomes (1976). More recent literature find that children from larger families have lower average education levels. Black et al. (2005) use a rich dataset on the entire population of Norway and find a negative correlation between family size and children's education. Silles (2010) finds that sibling size has an adverse causal effect on test scores and behavioural development. Li (2007) finds that student who is the single child or has only one sibling in the family is more likely to be enrolled in elite universities in China. Thus, family environment especially sibling size plays a prominent role in shaping the educational expectations of children.

Second, the effect of siblings on education returns is different for males and females, since children's educational attainment is the result of family decision making, in which the household balances efficiency and equity issues to determine the optimal distribution of educational resources (Becker, 1991). Researchers find patterns of discrimination in favor of boys in the allocation of household resources, including nutrients, health care, and education (Behrman, 1988; Deaton, 1989) . Parish and Wills (1992) verify that a brother reduces siblings' education in Taiwan. Butcher and Case (1994) examine the effect of the number and sex composition of a boy or girl's siblings on that children's educational attainment. They find that in the United States between 1920 and 1965, women's educational choices have been systematically affected by the sex composition of her siblings, and that men's choices have not. Hence, a child's education may be affected by the size and sex composition of siblings if the family faces loaning constraints, especially in the developing countries such as China.

Sons and daughters have different earnings potentials during the period in which they contribute to household income. Women historically spent less time in the labour force than men and for this reason the returns to education investment may have been lower for daughters than for sons. If sons receive a higher return to each level of schooling, we should expect to see not only sons receive more education, but also that the presence of sons reduces the educational attainment of daughters. Therefore, a girl with only sisters would receive more education than a girl with brothers, under the assumption of loaning constraints and not upon the exact shape of the parents' utility function (Butcher and Case, 1994). We will check this point in Chinese families.

Third, we follow literature and argue that education is a kind of public good affected by local investment on public facilities. Berik et al. (2007) also point out that privatization of healthcare and education disproportionately impedes women's and girls' access to these services. Women's increased care burden causes them to seek their children's help in household tasks, often cutting short their daughters' education. Even though good education facilities are located in large urban areas and provincial capitals, their locations are determined by historical factors, geographic characteristics, fiscal policy, political considerations, defense goals, and the like (Fleisher et al. (2010). Pal (2010) verifies that access to public infrastructure plays a crucial role on the presence of private schools in a community in India. He highlights the role of village-level public infrastructure (as reflected in the access to concrete road, electricity, piped water, phone and post office). Thus, it is reasonable to assume that public facilities tend to generate exogenous impacts on people’s earnings through education rather than that their locations have been the result of earnings of people who live there.

Maimaiti and Siebert (2009) investigate girls' school dropout rates in rural China and argue that a girl's education suffers when her larger water need for female hygiene purposes after menarche is not met because her household has poor access to water. Using the same data source as our paper, they find that menarche is associated with an increase in the school dropout rate, and indeed the effect is weaker for girls who have good access to water. Hence, they conclude that water engineering can thus contribute significantly to reducing gender education gaps in rural areas.

However, Oster and Thornton (2011) address sanitary product provision and argue that differences in dropout rate by girls with and without their period (or between girls and boys) in China cannot be explained by access to water without good attendance data as in Maimaiti and Siebert (2009), which is difficult to rule out omitted variable stories. In other words, access to tap water could be a proxy of public facility investment by local government which affects education of both males and females (but may affect girls more than boys as we see below), and possibly irrelevant to girls’ period. Thus, we only use access to tap water as a potential instrumental variable for education of both males and females, rather than an explanatory factor for gender education gap.

Last but not least, because we have employees from many different cohorts, individuals are in different sets of years and ages. The aggregated estimates of regressions may be mix-ups of many heterogeneous cohorts hence suffer serious composition biases (Solon et al., 1994). Therefore, to maximize efficiency, we need control for cohort effects. We estimate the regressions by separated gender and age cohorts to gain more understanding on patterns of education returns in China. The three age cohorts are people born before 1962, during 1962-1980 and born after 1980. The cut-off time choice of cohorts is based on the widely accepted structural break points in Chinese modern history to allow heterogeneous human capital accumulation in our study. The first structural break point is the Chinese famine during 1959-1961 resulted in the premature death of 30 million people (Lin 1990). Second, 1980 is considered a watershed year because of "One-Child Policy" and "reform and open-door" policy enacted at the end of the 1970s (Kang and Peng, 2013). Hence, the first age cohort includes people born before 1962; the second age cohort covers the largest baby boom periods (1962-1980); the youngest cohort includes people born after 1980, that is, the "One-Child Policy" cohort.

Overall, we adopt an IV approach and using the "numbers of brothers", "numbers of sisters" and "access to piped / tap water in house or courtyard" variables as instruments. The following two-step model describes the natural logarithm of hourly wage and years of schooling are normally applied to cope with the endogeneity of schooling:

$$(2) \qquad \ln W_i = \alpha X_i^{'} + \beta YS_i + \mu_i$$

$$(3) \qquad YS_i = \delta Z_i' + \nu$$

where Z_i denotes the vector of observed instrumental variables for years of schooling YS_i. Number of brothers (*Bro_i*), number of sisters (*Sis_i*), access to tap water (*$Water_i$*), and the other exogenous variables such as experience and location dummies are in the vector X_i, same as in the OLS regression. We estimate this model by gender and age cohorts. Thus, our examination of the gendered education returns associated with China's reforms begins with an overview of insights from structural adjustment programs since the end of 1980s and the economic reforms in transition economies in the 1990s and 2000s.

III. Data

The dataset used in this paper is the eight waves of the China Health and Nutrition Survey (CHNS, 1989, 1991, 1993, 1997, 2000, 2004, 2006 and 2009). The survey employs a multistage random-cluster sampling process to draw households from nine administrative divisions (Guangxi, Guizhou, Heilongjiang, Henan, Hubei, Hunan, Jiangsu, Liaoning, and Shandong) which include the developed coast and the inland. Four counties and 167 communities are randomly selected in each province. There are about 4,400 households with a total of 16,000 individuals in each survey (Liu et al. 2010, Li and Wu 2011). We use all eight waves to compare returns to education over time. The eight sample years represent distinct phases of economic reform and business cycle in China. Specifically, the year 1989, 1991 and 1993 represent the early stage of reform and transition, during which the Chinese leadership still sought reform within the socialist system and ‘‘a reform with no losers’’ (Lau et al., 2001); the year 1997, 2000 and 2004 represent the reloaded reform after Deng Xiaoping’s South China Tour in 1992, as the central government formally endorsed the full-fledged march towards market-orientated economy and accelerated steps toward integration with the global economy (Berik et al., 2007); 2006 and 2009 reflect the boom and bust during the most recent financial crisis.

Our sample only includes employees (for males aged 16-60 and females aged 16-52)[3] with positive annual gross income, working in sectors excluding “Agriculture, forestry, animal husbandry, fishery and water conservancy”. Self-employed and owners of private or individual enterprises have been excluded, because it is difficult to separate their wages from the profit income. Observations with missing values on years of schooling and experience have been dropped. Nominal hourly wage is the nominal annual earnings (include regular wages, bonuses, all kinds of subsidies and in-kind wages from the work unit) divided by annual working hours. We deflate nominal hourly earnings into real hourly earnings using provincial urban/rural Consumer Price Index based on 1995, provided by the National Bureau of Statistics.

In the CHNS dataset, the numbers of siblings are only surveyed on ever-married women under age 53 and their husbands in 2000, 2004, 2006 and 2009. Because the CHNS 1989-2009 are actually panel data, we can assume the sibling numbers are

[3] In China, the regulated retirement age is 60 for males and 55 for females. The CHNS dataset provides the sibling information only for ever-married females under age 53, so we use the cut-off age of 53 for our female sample.

invariant over time for all adult respondents and trace back their information of sibling numbers in other waves for those who have once appeared in the 2000-2009 waves. [4]

Table 1 shows the data description of the main variables over the period 1989-2009 by gender. Given the typically long working hours, the real hourly wage rates of both males and females were very low in China before 1997 (< 2 Yuan per hour), which jumped to about 3-4 Yuan per hour in 1997, and then have been rising to 10.19 Yuan for males and 7.49 Yuan for females in 2009. With respect to gender earnings differentials, males' wages have grown faster than females' over last two decades. The earnings ratio of women to men decreased mildly, from 0.83 in 1989 to 0.74 in 2009, which is consistent with Gustafsson and Li's (2000) findings.

The average years of schooling have increased from 8.39 years to 10.49 years over the entire period for males. The average years of schooling of females was 7.82 years in 1989, being less than males', but they caught up with males very quickly and surpassed them in recent years. The average numbers of brothers or sisters were less than two, suggesting that one family has 4 or 5 children which were declining slightly over time as the "One-Child Policy" was enacted after 1980. The percentages of

---

[4] For those missing values, we impute their sibling numbers using prediction function as follows: $Sib_i = f(B_i, Pr_i, Urban_i, Male_i)$, where Sib_i is the number of brothers (Bro_i) or number of sisters (Sis_i), B_i is the birth year, Pr_i is the vector of province dummies; $Urban_i$ is the urban dummy (1=urban, 0=rural), $Male_i$ is the gender dummy (1=male, 0=female).

households which had access to piped/tap water in house or courtyard had been increasing from about 70 percent in 1989 to about 90 percent in 2009. The average experience of male employees was around 21 years before 2004 and more than 24 years after then, partly reflecting the population ageing resulting from the "One-Child Policy". Female employees had less experience because they retired earlier than males. Moreover, less than half of the employee sample was from urban area. The oldest male cohort (born before 1962) decreased from 62 percent to 32 percent of our employee sample due to retirement, and the baby boom cohort after Chinese famine (born during 1962-1980) was above 50 percent after 1997. The "One-Child Policy" cohort (after 1980) entered the labour market in 2000, and comprised 14 percent of full sample in 2009. Female cohorts have very similar pattern to male cohorts.

(Table 1 is around here)

Table 2 describes the variables by gender and age cohorts in the pooled dataset of all eight waves. As expected, the youngest cohort has the highest real hourly wage rates of 6.8 Yuan per hour for males and 5.92 Yuan per hour for females. They also have the longest years of schooling of 10.59 years for males and 11.19 years for females, as well as much shorter experience. And, they have much less brothers and sisters than older cohorts for both genders which is consistent with the enforcement of "One- Child Policy" after 1980.

Although the "One-Child Policy" was meant to cover the country as a whole, people in rural areas have always allowed a second child if the first birth is a girl. In some rural areas and over some periods even three children were allowed. In addition, ethnical minority people who are subject to much looser restrictions have more opt-out from the policy (Peng, 1991, Cameron 2011). In the urban areas, however, the policy has been strictly enforced since it was introduced. Those who obey the policy are financially rewarded while those who violate the policy are subject to fines and their children face higher fees for accessing education and health services. In some cases, children are denied these services (Kane an Choi, 1999; Zhang and Sturm, 1994). Hence, once a family had a boy, most parents would stop seeking legal or illegal way to have another child because either way is costly.

(Table 2 is around here)

Under the restriction of policy, Chinese parents would like to seek a way to have another child if they had the first birth of girl. Once they had a boy, they would stop because of high cost of fine. Hence, parental sex selection on children may distort the sex composition of siblings. It is not surprising that in the "One-Child Policy" cohort, males have fewer brothers (0.43) than sisters (0.81), while females have more brothers (0.82) than sisters (0.64). It is a prominent policy distortion of sex composition of siblings, compared with older cohorts who have broadly similar brother and sister numbers for both males and females (around 2 for the cohort born before 1962, around 1.6 for the cohort born 1962-1980). It can also explain the less proportions of urban residence in younger cohorts as people in rural areas have opt-out from the policy. Thus, gender and age cohorts show much heterogeneous characteristics of variables in our sample, supporting estimation by gender and age cohorts.

IV. Empirical Results

A lot of literature warn about the endogeneity problem of years of schooling and suggest the instrumental variables for unbiased estimations. We apply a 2 Stage Least Square (2SLS) procedure for estimation. In the first step, we regress years of schooling on instruments and get the 1SLS estimator by cohort using equation (3). Then we calculate the predicted values of years of schooling for each cohort. Hence, the variable of years of schooling can be decomposed into two components: a linear combination of instruments (predicted years of schooling) and a random component. In the second step, we estimate equation (2) using the predicted years of schooling. And, we apply this GIV method for estimation year by year and cohort by cohort to overcome the possible over-identification and heteroskedasticity problem.[5]

A. First step of GIV

Results of the first step of GIV by gender and age cohorts are reported in Table 3.

[5] See Card (2001)for a theoretical treatment of the interpretation of instrumented variables and the practical application in STATA (Baum et al., 2003).

Estimation would perform poorly when instruments were weak, so we need test the validity of the instruments (number of brothers, number of sisters, and access to tap water). We report the Cragg-Donald Wald F statistics (Cragg and Donald, 1993), critical values of which have been compiled by Stock and Yogo (2002). We use these critical values to judge the validity of instruments, and find that linear combination of instruments fit years of schooling very well except the male cohort born after 1980.

(Table 3 is around here)

Moreover, we stack the cohort-specific regressions using the Chi2 statistics in the Seemingly Unrelated Estimation (*SUEST)* (Weesie, 1999) to test whether coefficients of these IVs significantly different from zero. The Chi-2 statistics are 308.14 for the combined two regressions of the full samples of males and females; 274.4 for the combined regressions of three male age cohorts; and 240.39 for the combination of regressions of three female age cohorts. The Chi2 tests easily reject the hypothesis that the three IVs are not significantly associated with years of schooling. Hence, we regard three IVs as valid instrumental variables for years of schooling.

In Table 3, with constraint of economic resources for education investment, parents need allocate intrahousehold resource on children’s education. We find that the more does the respondent have siblings the less are his/her years of schooling. For the full sample of males, one more brother would decrease a boy’s schooling by about 0.16 year. In the similar way, one more sibling (brother or sister) would decrease a girl's schooling by about 0.17 year for the full sample of females. However, boys' education is much less affected by their sisters, as one more sister would decrease a boy's schooling only by about 0.06 year. It reflects the long tradition of parental preference for sons and less education investment on daughters in China. In a patrilineal and patrilocal society like China, sons and their wives are expected to live with sons' parents; daughters are married out and become part of another family; children are named with their father's last name, not the mother's. Therefore, only sons who can guarantee the provision of financial support and care for the old parents could inherit the family name and carry on the family line (Li and Wu, 2011).[6] In

---

[6] The parental son preference in China could be traced back to the origins of ancestral worship in the second and third millennia B.C. The patrilocal and patrilineal familial system developed during imperial state reinforced this preference(Lee and Wang, 1999) (Bray, 1997) Li and Wu 2011 . In one of the oldest poems in China, sons and daughters were treated in different ways in family (sleeping places, clothes and toys for education):

Sons shall be born to him : --

They will be put to sleep on couches ;

They will be clothed in robes ;

They will have sceptres to play with ;……

Daughters shall be born to him : --

They will be put to sleep on the ground ;

They will be clothed with wrappers ;

such a society, girls would have been allocated less resource for education than their brothers by parents.

For different age cohorts, we find that the negative effects of siblings are more significant in two older age cohorts (born before 1980) than in the youngest "One-Child Policy" cohort after 1980. Especially in the cohort born during 1962-1980, the resources constraints are more serious after 1959-1961 famine and thereafter baby boom, suggesting more careful relocation of intrahousehold resource. The baby boom cohort is the current majority of labour force so dominates the significantly negative effects of siblings (brothers or sisters) on years of schooling in the full sample.

Moreover, according to the opt-out articles of the "One-Child Policy", only ethnical minority and people in rural areas with the first birth of a girl are allowed to have more than one child. Once a family had a boy, parents would stop seeking legal or illegal way to have another child because either way is costly. Hence, the preference of first birth of boy is strengthened by the policy and finally affects the education investment intrahousehold. In the "One-Child Policy" cohort, it is not surprising that education of males is not significantly affected by siblings (brothers or sisters). For the females, their education would not be affected by their brothers. It is possibly because there are more intrahousehold resources as well as better public

---

They will have tiles to play with .

(***Shi Jing [Book of Odes]***, translated by Legge 1876, University of Virginia Library, http://etext.lib.virginia.edu/chinese/).

facilities outside available after economic reforms launched in 1980. The tuition fees are removed off for the 9-year compulsory education with the compulsory policy launched in 1986. Thus, intrahousehold resource allocation for children's education is much easier after 1980 than before.

However, the education of a girl would be heavily affected (-0.379) if she had another sister, suggesting their parents have no legal chance to have the third child for a son. Having no male descendant would bring structural breaks in family decisions on consumption, saving and children education investment. Taking into account the influence of Chinese traditional cultural norms on women's marriage and care responsibilities, the responsibility of caring old parents is mainly upon their sons and daughters-in-law (Liu et al., 2000). The parents without male descendants may expect no financial support and care from daughters in their old ages and would not reduce their current consumption for daughters' education investment. In other words, parents without son may save more money from reduction of their daughters' education to help their older life in case of no financial support and care from daughters.

At the same time, under the stricter policy, those parents choose to, legally or illegally have another child after the first birth of a girl are more likely to prefer sons and large family size. They are a selected subgroup of parents of sample in the cohort born after 1980. Hence, they are also more disappointed with the second girl and relocate less intrahousehold resource on children's education. Comparing with their spouses, women are found to be more likely to spend family resources on nutrition, education, and health-related commodities (Von Braun, 1988; Thomas, 1990), so play an important role on intrahousehold resource allocation on children education (especially for daughters). However, Li and Wu (2011) find that the mothers with first birth of girl would have less bargaining power on intrahousehold resource allocation and have less resource for children's education. This result is consistent with the view of feminist scholars that social and cultural norms are important determinants of women's education. We do not find this kind of distortion in older cohorts, because females with sisters do not mean that they have no brother before the "One-Child Policy". Thus, as feminist economists argued, the seemingly gender neutral policies implemented through social and cultural norms that bear and transmit gender biases can impede gender equity, and that gender inequality, in turn, can hinder the attainment of other human development objectives (Seguino and Grown, 2006; Ding et al., 2009).

Access to tap water improves the years of schooling for the three cohorts and both genders, which has significantly positive effects on children's education especially for girls. If a household has access to tap water, implying better public facilities provided by local government during the period of children's education, the years of schooling of boys (full sample) would increase by 1.5 years and even more for girls by 1.7 year. Hence, education of females is more dependent on outside public goods than males. For different age cohorts, we find the effects of access to water are less important for younger cohorts than for older cohorts. For example, access to water improves years of schooling of males by 1.5 year (females by 1.9 year) in the cohorts born before 1962, compared with by 1.1 year for males (females by 1.5 year) for in the cohorts born after 1980. Thus, the external public goods are less important for family decision on education over time, as more and more private resources are available intrahousehold.

B. ***OLS vs. second step of GIV***

Table 4 compares OLS and GIV estimation year by year. Panel A shows the basic results of OLS estimation using equation (1). In the full sample, females' returns for an additional year of schooling (average 5.4 percent) are higher than males' (average 3.7 percent). The earnings returns for an additional year of schooling increase from about 2 percent in 1989 for both genders to 7 percent for males and 10.3 percent for females in 2009. However, for both males and females, we see a decline of education returns after 2004. The big expansion of higher education happened in 1999, as the enrolment of higher education has increased 4 times in 7 years (Qian and Smyth 2011). The decline of education returns after 2004 just reflect the graduates entering the labour market.

(Table 4 is around here)

Panel B in Table 4 shows the estimation results of the second step of GIV regressions by year and gender using equation (2) and predicted variable of years of schooling. The returns to schooling are insignificant for males in 1991, 1993 and 1997 and for females in 1993. Before 2004, education returns for females are higher than males as we find in OLS results. However, by eliminating the endogeneity problem, males' returns to schooling are higher than their OLS estimators over the entire period. Females' returns to schooling are also higher than their OLS estimators before 2004, but lower than their OLS estimators after 2000. Education returns for females are lower than males after 2000, which are different from OLS results. Thus, there are downwards biases of OLS estimation for males in all years as Card (1999) and Li and Luo (2004) argued. For females, however, downwards biases of OLS estimation are only for data before 2000. For females after 2000, the OLS estimation has upwards biases on education returns as (Chen and Hamori, 2009) argued.

These results show the necessity of instruments of siblings and tap water, and prove the over-optimistic assessment of females' education returns after 2000 may be from the ignorance of negative effects of endogeneity problem. In other words, if females with higher ability have more education, OLS results would give over-optimistic judgment on females' education returns. However, we do not know what forces are behind the drop of females' education returns after 2000 and what take account for the gap of education returns between males and females. Hence, we push the study further by estimating the model by gender and age cohorts.

Table 5 compares OLS and GIV estimation cohort by cohort. Panel A illustrates the interactive variables of year of schooling and year dummies for the pooled data by gender and age cohort. For the full sample of both genders, the returns to schooling are quite stable around 2 percent for an additional year of schooling during the period of 1989-1997 and then jump to 4.7 (=1.5+3.2) percent for males and 6 (=2.5+3.5) percent for females in 2000,[7] then nearly double in 2004. In 2009, the returns to schooling are 7.6 (=1.5+6.1) percent for males and 10.7 (=2.5+8.2) percent for females. It is consistent with what we find in Table 4 that females have always higher education returns than males.

For different age cohorts, education returns are very similar to the full sample except the cohort born after 1980. Because this cohort entered the labour force gradually after 2000, they experienced the higher education expansion and brought "newly" accumulated human capital into the labour force. For males, the education returns of the "One-child Policy" cohort have a dramatically jump from about 8.2 percent in 2004 for an additional year of schooling to 16.2 (=8.2+8) percent in 2006 and 15.1 (=8.2+6.9) percent in 2009. For females, the education returns are even higher which increase from about 10.3 percent for an additional year of schooling in 2004 to 19.8 (=10.3+9.5) percent in 2006, and 19.1 (=10.3+8.8) percent in 2009. Hence, return rates for the youngest cohort are much higher than the older cohorts. It suggests that the human capital accumulated during the "pre-transition" period of economy may have an increasing returns in the "early-transition", but would have less returns in the "late transition" than the "new" human capital accumulated during the "early-transition" period, supporting the findings of Flabbi et al. (2008) and Kang and Peng (2013).

---

[7] The table in Appendix shows the aggregated effects of each year (= the baseline coefficient + incremental effects of year dummies) in Table 5.

(Table 5 is around here)

Panel B shows the returns to schooling by cohort and gender using GIV. GIV estimation for returns to schooling is all higher than OLS results except females of older cohorts during the period of 2004-2009, which is consistent what we find in Table 4. It confirms that the biases of OLS are mainly downwards except upwards biases for females (especially in old cohorts) in recent years. And, we find that the "One-Child Policy" cohort still shows the highest education returns in GIV, and the most serious downwards biases of OLS. For example, the males born after 1980 have 11.8 percent returns for an addition year of schooling in 2004, and then 24.3 (=11.8+12.5) percent in 2006 and 26.4 (=11.8+14.6) percent in 2009. The biases between OLS and GIV estimation are about 3−11 percent for males.

The females born after 1980 have 17.5 percent returns for an addition year of schooling in 2004, and then 33.3 (=17.5+15.8) percent in 2006 and 34.2 (=17.5+16.7) percent in 2009, which are higher than males and corresponding OLS results but still consistent with Card (1999) and Li and Luo (2004). The biases between OLS and GIV estimation are about 7−15 percent for females.

Therefore, education returns of older cohorts have dropped in recent years for both genders, especially for females. The young cohort has much higher education returns than older cohorts maybe because of the accumulation of "new" human capital after the "reform and open-door" policy. As an increasing proportion of labour force, the young females are helping stop the reversal of education returns gap between females and males. Upwards biases of OLS only happen in females of older cohorts after 2004. It suggests that only senior female workers with higher ability who also have more education, still competed with the younger cohorts with more higher education, even though their human capital accumulated from the education during the pre-transition has been substituted by “new” human capital accumulated by the younger cohort.

V. Conclusions

We examine the returns to schooling in China using the China Health and Nutrition Survey (CHNS) dataset. In the baseline Ordinary Least Square (OLS) regression, the returns to one more year of schooling increase from about 2 percent in 1989 for both genders to 7 percent in 2009 for males, and 10.3 percent for females. The returns to schooling for females in all three cohorts are higher than males over the entire period, which is consistent with the literature using OLS for studies of Chinese education returns.

However, we find more interesting results after overcoming the well-known endogeneity problem of education. We use the Generalized Instrumental Variable (GIV) regressions with three instruments: number of brothers, number of sisters and access to tap water which show strong associations with children’s education, especially for girls. As far as siblings are concerned, one more brother would decrease a boy’s schooling by about 0.16 year. One more sibling (brother or sister) would decrease a girl’s schooling by about 0.17 year. As expected from the long tradition of son preference culture, boys’ education is much less affected by their sisters, as one more sister would decrease a boy’s schooling only by about 0.06 year. Access to tap water, as a proxy of public facilities, always has significantly positive effect on children’s education especially for girls. Education of females is more dependent on outside public goods than males.

Comparing the OLS and GIV estimations for returns to schooling year by year, we find that there are downwards biases of OLS estimations for males in all years. However, for females, downwards biases of OLS estimation are only for data before 2004. Comparing OLS and GIV estimations cohort by cohort can provide more detailed information. Regressions with year dummies are all higher than OLS results except older female during the period of 2004-2009. It confirms that the biases of OLS are mainly the downwards biases except upwards biases for females in recent years.

Therefore, we find that the “One-Child-Policy” cohort shows the highest education returns in GIV and the most serious downwards biases from OLS. In this cohort, females have higher education returns than males’ and corresponding OLS results. These return rates for the youngest cohort are much higher than the older cohorts after 2004. It suggests that the human capital accumulated during the “pre-transition” period of economy may have an increasing returns in the “early-transition”, but would have less returns in the “late-transition” than the “new” human capital accumulated during the transition period, verifying the findings in

Flabbi et al. (2008) and Kang and Peng (2013). However, since the “One-Child Policy” cohort has much fewer observations than other cohorts in our sample, any formal interpretation should be concerned with caveats and demand further research.

Table 1: Data description by gender and year

Male Employees	All	1989	1991	1993	1997	2000	2004	2006	2009
Real hourly wage rate (Yuan)	4.23	1.27	1.53	1.91	2.99	5.82	6.42	7.51	10.19
Years of schooling (Years)	9.54	8.39	8.81	9.12	9.48	9.91	10.43	10.67	10.49
Number of brothers	1.79	1.89	1.89	1.84	1.80	1.79	1.77	1.81	1.63
Number of sisters	1.75	1.81	1.86	1.81	1.80	1.68	1.72	1.72	1.67
Access to tap water	0.82	0.74	0.79	0.81	0.83	0.82	0.87	0.89	0.91
Experience (Years)	22.26	20.39	21.97	21.36	21.54	20.88	23.82	24.63	25.19
Urban	0.45	0.47	0.48	0.44	0.45	0.41	0.47	0.46	0.44
Cohort (born before 1962)	0.51	0.62	0.66	0.59	0.49	0.42	0.42	0.40	0.32
Cohort (born 1962-1980)	0.45	0.38	0.34	0.41	0.50	0.53	0.52	0.51	0.54
Cohort (born after 1980)	0.04					0.05	0.07	0.09	0.14
Female Employees	All	1989	1991	1993	1997	2000	2004	2006	2009
Real hourly wage rate (Yuan)	3.50	1.15	1.37	1.49	4.19	4.09	5.60	6.00	7.49
Years of schooling (Years)	9.40	7.82	8.72	8.78	9.31	9.86	10.67	10.78	10.93
Number of brothers	1.73	1.89	1.81	1.83	1.76	1.76	1.70	1.71	1.50
Number of sisters	1.80	2.05	1.92	1.90	1.80	1.76	1.74	1.72	1.66
Access to tap water	0.86	0.77	0.85	0.85	0.86	0.85	0.89	0.93	0.93
Experience (Years)	18.42	18.13	17.70	18.00	17.47	17.50	19.08	20.18	20.33
Urban	0.51	0.51	0.54	0.50	0.52	0.47	0.53	0.49	0.50
Cohort (born before 1962)	0.38	0.54	0.53	0.48	0.35	0.30	0.26	0.21	0.11
Cohort (born 1962-1980)	0.57	0.46	0.47	0.52	0.65	0.62	0.62	0.66	0.68
Cohort (born after 1980)	0.05					0.07	0.12	0.13	0.21

Data source: the CHNS dataset 1989-2009

Table 2: Data description by gender and age cohort

Male Employees	Born before 1962	Born 1962-1980	Born after 1980
Real hourly wage rate (Yuan)	3.77	4.62	6.80
Years of schooling (Years)	8.96	10.12	10.59
Number of brothers	2.07	1.58	0.43
Number of sisters	1.89	1.65	0.81
Access to tap water	0.83	0.81	0.82
Experience (Years)	30.86	13.73	5.45
Urban	0.49	0.41	0.38
Female Employees	**Born before 1962**	**Born 1962-1980**	**Born after 1980**
Real hourly wage rate (Yuan)	3.17	3.53	5.92
Years of schooling (Years)	8.35	9.98	11.18
Number of brothers	2.01	1.58	0.82
Number of sisters	2.06	1.67	0.64
Access to tap water	0.89	0.84	0.79
Experience (Years)	27.32	13.36	4.77
Urban	0.58	0.47	0.43

Data source: the CHNS dataset 1989-2009

Table 3: First step of GIV regressions

Male Employees	All	Born before 1962	Born 1962-1980	Born after 1980
Number of brothers	-0.163***	-0.199***	-0.131***	-0.325
	0.026	***0.036***	***0.032***	***0.278***
Number of sisters	-0.060**	0.025	-0.175***	0.054
	0.026	***0.034***	***0.041***	***0.203***
Access to tap water	1.492***	1.512***	1.417***	1.056***
	0.108	***0.147***	***0.113***	***0.292***
R-squared	0.321	0.389	0.157	0.443
N	11415	5955	5020	440
Cragg-Donald Wald F-test	78.46***	44.81***	62.67***	5.19
Female Employees	**All**	**Born before 1962**	**Born 1962-1980**	**Born after 1980**
Number of brothers	-0.167***	-0.066	-0.214***	-0.011
	0.043	***0.062***	***0.058***	***0.262***
Number of sisters	-0.174***	-0.123***	-0.202***	-0.379*
	0.031	***0.047***	***0.041***	***0.214***
Access to tap water	1.742***	1.867***	1.641***	1.534***
	0.135	***0.23***	***0.137***	***0.288***
R-squared	0.396	0.501	0.201	0.475
N	7926	3176	4309	441
Cragg-Donald Wald F-test	68.32***	23.32***	60.18***	13.44***

Note: The *italic* standard errors are adjusted by clusters. The significant levels are * for 10%; ** for 5% and *** for 1%. Experience, experience square, urban dummy, provincial dummies and year dummies are not reported. Critical values complied by Stock and Yogo (2002) for the Cragg-Donald F statistic are around 10 for our estimations, which are passed easily except the male cohort born after 1980.

Table 4: Education returns estimation using OLS and GIV regressions (by year and gender)

	All	1989	1991	1993	1997	2000	2004	2006	2009
Panel A. OLS regression by year, males									
Year of schooling	0.037***	0.026***	0.008	0.01	0.025***	0.044***	0.083***	0.081***	0.070***
	0.003	***0.005***	***0.005***	***0.007***	***0.008***	***0.01***	***0.009***	***0.008***	***0.008***
R-squared	0.596	0.135	0.101	0.065	0.08	0.072	0.184	0.206	0.127
N	9961	1560	1577	1330	1210	1124	973	1045	1142
OLS regression by year, females									
Year of schooling	0.054***	0.022***	0.028***	0.018**	0.037***	0.056***	0.111***	0.108***	0.103***
	0.004	***0.007***	***0.006***	***0.008***	***0.009***	***0.012***	***0.014***	***0.011***	***0.01***
R-squared	0.593	0.093	0.089	0.107	0.087	0.081	0.268	0.244	0.286
N	6708	1029	1106	909	832	718	677	686	751
Panel B. Second step of GIV regression by year, males									
Schooling	0.070***	0.056***	0.011	0.047	0.032	0.045*	0.134***	0.116***	0.110***
	0.008	***0.02***	***0.024***	***0.031***	***0.026***	***0.026***	***0.023***	***0.018***	***0.021***
R-squared	0.591	0.115	0.1	0.064	0.069	0.046	0.118	0.139	0.082
N	9607	1462	1510	1270	1150	1086	963	1030	1136
Second step of GIV regression by year, females									
Schooling	0.062***	0.057**	0.067***	0.035	0.047*	0.075***	0.096***	0.055**	0.089***
	0.009	***0.022***	***0.023***	***0.026***	***0.025***	***0.024***	***0.027***	***0.023***	***0.024***
R-squared	0.581	0.09	0.081	0.105	0.069	0.056	0.145	0.085	0.166
N	6569	997	1084	889	809	700	669	678	743

Note: "Schooling" in the second step of GIV regression is the predicted value from the first step in Table 3. The *italic* standard error adjusted by clusters. The significant levels are * for 10%; ** for 5% and *** for 1%. Provincial dummy variables and year dummies are not reported.

Table 5: OLS and GIV regressions by gender and age cohort

	All	Born before 1962	Born 1962-1980	Born after 1980
Panel A. OLS regression by age cohort, males				
Schooling	0.015***	0.021***	0.028**	
	0.004	***0.004***	***0.013***	
Schooling*year1991	-0.010*	-0.010*	-0.025	
	0.006	***0.006***	***0.017***	
Schooling*year1993	-0.004	-0.005	-0.019	
	0.007	***0.008***	***0.018***	
Schooling*year1997	0.01	0.011	-0.009	
	0.009	***0.01***	***0.017***	
Schooling*year2000	0.032***	0.041***	0.005	
	0.009	***0.012***	***0.017***	
Schooling*year2004	0.067***	0.068***	0.058***	0.082**
	0.009	***0.012***	***0.017***	***0.033***
Schooling*year2006	0.074***	0.077***	0.059***	0.080***
	0.008	***0.011***	***0.017***	***0.028***
Schooling*year2009	0.061***	0.073***	0.043***	0.069***
	0.008	***0.016***	***0.017***	***0.025***
R-squared	0.607	0.594	0.616	0.301
N	9961	5383	4255	323
OLS regression by age cohort, females				
Schooling	0.025***	0.036***	0	
	0.006	***0.007***	***0.01***	
Schooling*year1991	0	-0.006	0.022	
	0.008	***0.009***	***0.014***	
Schooling*year1993	-0.009	-0.013	0.003	
	0.009	***0.011***	***0.018***	
Schooling*year1997	0.006	-0.004	0.031*	
	0.011	***0.015***	***0.016***	
Schooling*year2000	0.035***	0.049***	0.035**	
	0.011	***0.014***	***0.015***	
Schooling*year2004	0.086***	0.071***	0.107***	0.103**
	0.011	***0.02***	***0.015***	***0.04***
Schooling*year2006	0.073***	0.046***	0.098***	0.095***
	0.01	***0.014***	***0.016***	***0.023***
Schooling*year2009	0.082***	0.031	0.114***	0.088***
	0.01	***0.021***	***0.015***	***0.019***
R-squared	0.608	0.57	0.609	0.347
N	6708	2742	3636	330
Panel B. Second step of GIV regressions by age cohort, males				
Schooling	0.031***	0.061***	0.06	
	0.011	***0.016***	***0.039***	

Schooling*year1991	-0.013	-0.014	-0.068	
	0.011	***0.011***	***0.052***	
Schooling*year1993	0.014	0.009	-0.021	
	0.014	***0.015***	***0.062***	
Schooling*year1997	0.003	-0.006	-0.074	
	0.015	***0.016***	***0.051***	
Schooling*year2000	0.039***	0.041**	-0.032	
	0.014	***0.017***	***0.051***	
Schooling*year2004	0.072***	0.073***	0.103**	0.118*
	0.015	***0.021***	***0.048***	***0.064***
Schooling*year2006	0.098***	0.088***	0.130***	0.125***
	0.013	***0.018***	***0.049***	***0.048***
Schooling*year2009	0.079***	0.074***	0.046	0.146***
	0.015	***0.022***	***0.049***	***0.055***
R-squared	0.596	0.582	0.608	0.278
N	9607	5212	4082	313
Second step of GIV regressions by age cohort, females				
Schooling	0.048***	0.068***	0.077***	
	0.015	***0.021***	***0.026***	
Schooling*year1991	-0.014	-0.015	-0.017	
	0.015	***0.017***	***0.04***	
Schooling*year1993	-0.023	-0.021	-0.076*	
	0.016	***0.017***	***0.041***	
Schooling*year1997	-0.022	-0.025	-0.067*	
	0.018	***0.02***	***0.039***	
Schooling*year2000	0.024	0.034	-0.026	
	0.017	***0.021***	***0.035***	
Schooling*year2004	0.052***	0.034	0.068**	0.175***
	0.018	***0.026***	***0.034***	***0.05***
Schooling*year2006	0.031*	0.003	0.049	0.158***
	0.018	***0.024***	***0.037***	***0.044***
Schooling*year2009	0.075***	0.053*	0.070*	0.167***
	0.019	***0.028***	***0.037***	***0.043***
R-squared	0.586	0.567	0.584	0.339
N	6569	2704	3545	320

Note: "Schooling" is the predicted value from the first step in Table 3. The standard error adjusted by clusters. The *italic* standard error adjusted by clusters. The significant levels are * for 10%; ** for 5% and *** for 1%. Experience, experience square, urban dummy, provincial dummies variables are not reported. The significant levels are * for 10%; ** for 5% and *** for 1%. Provincial dummy variables and year dummies are not reported.

Appendix Coefficients table for years of schooling in Table 5

	All	Born before 1961	Born 1962-1980	Born after 1980
Panel A. OLS regression by age cohort, males				
1989	0.015	0.021		
1991	0.005	0.011	0	
1993	0.015	0.021	0	
1997	0.015	0.021	0	
2000	0.047	0.062	0.032	
2004	0.082	0.089	0.085	0.082
2006	0.089	0.098	0.087	0.162
2009	0.076	0.094	0.071	0.151
OLS regression by age cohort, females				
1989	0.025	0.036		
1991	0.025	0.036	0.022	
1993	0.025	0.036	0.022	
1997	0.025	0.036	0.053	
2000	0.06	0.085	0.057	
2004	0.111	0.107	0.129	0.103
2006	0.098	0.082	0.121	0.198
2009	0.107	0.036	0.136	0.191
Panel B. Second step of GIV regressions by age cohort, males				
1989	0.031	0.061		
1991	0.031	0.061	0	
1993	0.031	0.061	0	
1997	0.031	0.061	0	
2000	0.07	0.102	0	
2004	0.103	0.134	0.161	0.118
2006	0.129	0.149	0.188	0.243
2009	0.11	0.135	0.103	0.264
Second step of GIV regressions by age cohort, females				
1989	0.048	0.068		
1991	0.048	0.068	0	
1993	0.048	0.068	0	
1997	0.048	0.068	0	
2000	0.048	0.068	0	
2004	0.1	0.068	0.135	0.175
2006	0.079	0.068	0.116	0.333
2009	0.123	0.121	0.137	0.342

Notes: figures are the sum of the baseline returns and incremental effects of year dummies. Insignificant incremental effects are regarded as 0s.

Cameron, Lisa., Nisvan Erkal, Lata Gangadharan, and Xin Meng. （2011）. Little Emperors—Behavioral Impacts of China’s One-Child Policy, NBER working paper.

Kane, Penny and Choi, Ching, 1999, “China’s one child family policy”, *British Medical Journal*, 319(7215), pp. 992-994.

Peng, Xizhe. 1991, Demographic Transition in China: Fertility Trends since the 1950s (Studies on Contemporary China). Oxford: Clarendon Press.

Qian, Joanne Xiaolei and Russell Smyth. (2011). Educational expenditure in urban China: income effects, family characteristics and the demand for domestic and overseas education Applied Economics Volume 43, Issue 24, 2011. pages 3379-3394

Zhang, Junsen and Sturm, Roland, 1994, "When do couples sign the one-child certificate in urban China?", *Population Research and Policy Review,* 13 (1), pp. 69-81.

References

Baum, Christopher F., Schaffer, Mark E. and Stillman, Steven 2003. Instrumental variables and GMM: Estimation and testing. *The Stata Journal*.

Becker, Gary S. 1960. An Economic Analysis of Fertility. *NBER working paper*.

Becker, Gary S. 1991. A Treatise on the Family. *Cambridge, MA: Harvard University Press*.

Becker, Gary S. and Lewis, H. Gregg 1973. On the Interaction between the Quantity and Quality of Children. *Journal of Political Economy,* LXXXI**,** S279–S288.

Becker, Gary S. and Tomes, Nigel 1976. Child Endowments and the Quantity and Quality of Children. *Journal of Political Economy,* LXXXIV.

Behrman, Jere R. 1988. Intrahousehold Allocation of Nutrients in Rural India: Are Boys Favored? *Oxford Economic Papers,* XL**,** 32-54.

Belley, Philippe and Lochner, Lance 2007. The changing role of family income and ability in determining educational achievement. *Journal of Human Capital,* 1**,** 37-89.

Berik, Günseli, Dong, Xiao-Yuan and Summerfield, Gale 2007. China's Transition and Feminist Economics. *Feminist Economics,* 13**,** 1-33.

Bernhardt, Kathryn 1995. The Inheritance Rights of Daughters: the Song Anomaly? *Modern China* 21**,** 269–309.

Black, Sandra E., Devereux, Paul J. and Salvanes, Kjell G. 2005. The more the marrier? The effect of family size and birth order on children's education. *The Quarterly Journal of Economics*.

Blundell, Richard, Dearden, Lorraine and Sianesi, Barbara 2005. Evaluating the

Effect of Education on Earnings: Models, Methods and Results from the National Child Development Survey. *Journal of the Royal Statistical Society. Series A (Statistics in Society),* 168**,** 473-512.

Bray, Francesca 1997. Technology and Gender: Fabrics of Power in Late Imperial China. *Berkeley, Calif.: University of California Press*.

Butcher, Kristin F. and Case, Anne 1994. The Effect of Sibling Sex Composition on Women's Education and Earnings. *The Quarterly Journal of Economics,* 109.

Card, David 1999. The causal effect of education on earnings. *Handbook of Labor Economics,* Vol. 3A, Elsevier Science North-Holland, Amsterdam, New York and Oxford**,** 1801-1863.

Card, David 2001. Estimating the return to schooling: progress on some persistent econometric problems. *Econometrica,* 69**,** 1127-1160.

Chen, Guifu and Hamori, Shigeyuki 2009. Economic returns to schooling in urban China: OLS and the instrumental variables approach. *China Economic Review,* 20**,** 143-152.

Cragg, J. G. and Donald, S. G. 1993. Testing identifiability and specification in instrumental variable models. *Econometric Theory,* 9**,** 222-240.

Dearden, Lorraine 1999. The effects of families and ability on men's education and earnings in Britain. *Labour Economics,* 6**,** 551-567.

Deaton, Angus 1989. Looking for Boy-Girl Discrimination in Household Expenditure Data. *The World Bank Economic Review,* III**,** 1-15.

Ding, Sai, Dong, Xiao-Yuan and Li, Shi 2009. Women's Employment and Family Income Inequality during China's Economic Transition. *Feminist Economics,* 15**,** 163-190.

Edlund, Lena, Li, Hongbin, Yi, Junjian and Zhang, Junshen 2007. Sex Ratios and Crime: Evidence from China's One-Child Policy. *IZA Working Paper No. 3214*.

Flabbi, L., Paternostro, S. and Tiongson, E. 2008. Returns to education in the economic transition: A systematic assessment using comparable data. *Economics of Education Review,* 27**,** 724-740.

Fleisher, Li, Haizheng and Zhao, Min Qiang 2010. Human capital, economic growth, and regional inequality in China. *Journal of Development Economics,* 92**,** 215-231.

Fleisher, B. and Wang, X. 2005. Returns to schooling in China under planning and reform. *Journal of Comparative Economics,* 33**,** 265-277.

Freeman, Richard B. 2007. Labor Market Institutions around the World. *NBER working paper*.

Gal, Susan and Kligman, Gail 2000. Reproducing Gender: Politics, Publics, and Everyday Life After Socialism. *Princeton, NJ: Princeton University Press*.

Gustafsson, Bjo ̈Rn and Li, Shi 2000. Economic transformation and the gender earnings gap in urban China. *Journal of Population Economics,* 13**,** 305-329.

Harmon, C., Oosterbeek, Hessel and Walker, Ian 2004. The Returns to Education: Microeconomics. *Journal of Economic Surveys,* 17**,** 115-155.

Harmon, C. and Walker, I. 2000. Returns to the Quantity and Quality of Education: Evidence for Men in England and Wales. *Economica*, 67**,** 19-35.

Harmon, C., Walker, I. and Westergaard-Nielsen, N. 2001. Education and earnings in Europe: a cross country analysis of the returns to education. *Edward Elgar Publishing Limited*.

Heckman, James J. and Li, Xuesong 2004. Selection Bias, Comparative Advantage and Heterogeneous Returns to Education: Evidence from China in 2000. *Pacific Economic Review,* 9**,** 155-171.

Kang, L. and Peng, F. 2013. A Selection Analysis on Education Returns in China. *Post-Communist Economies*, 25.

Lau, Lawrence J., Qian, Yingyi Qian and Roland, Gerard 2001. Reform without Losers: An Interpretation of China's Dual-Track Approach to Transition. *Journal of Political Economy,* 101**,** 120–143.

Lee, James and Wang, Feng 1999. One Quarter of Humanity: Malthusian Mythology and Chinese Realities. *Harvard University Press, Cambridge*.

Li 2007. Family background, financial constraints and higher education attendance in China. *Economics of Education Review,* 26**,** 725-735.

Li, Haizheng and Luo, Yi 2004. Reporting errors, ability heterogeneity, and returns to schooling in China. *Pacific Economic Review,* 9**,** 191-207.

Li, Lixing and Wu, Xiaoyu 2011. Gender of Children, Bargaining Power, and Intrahousehold Resource Allocation in China. *The Journal of Human Resources,* 46**,** 295–316.

Liu, Pak-Wai, Meng, Xin and Zhang, Junsen 2000. Sectoral gender wage differentials and discrimination in the transitional Chinese economy. *Journal of Population Economics,* 13**,** 331-352.

Maimaiti, Yasheng and Siebert, W. S. 2009. The Gender Education Gap in China: The Power of Water. *IZA DP No. 4108*.

Meng, Xin and Kidd, Michael P. 1997. Labor Market Reform and the Changing Structure of Wage Determination in China’s State Sector during the 1980s. *Journal of Comparative Economics,* 25**,** 403-421.

Nee, Victor 1996. The Emergence of a Market Society: Changing Mechanisms of Stratification in China. *The American Journal of Sociology,* 101**,** 908-949.

Oster, Emily and Thornton, Rebecca 2011. Menstruation, Sanitary Products, and School Attendance: Evidence from a Randomized Evaluation *American Economic Journal: Applied Economics,* 3**,** 91-100.

Pal, Sarmistha 2010. Public infrastructure, location of private schools and primary school attainment in an emerging economy. *Economics of Education Review,* 29**,** 783-794.

Parish, William L. and Willis, Robert J. 1992. Daughters, Education and Family Budgets: Taiwan Experiences. *mimeo, University of Chicago, April*.

Peng, Yusheng 1992. Wage Determination in Rural and Urban China: A Comparison of Public and Private Industrial Sectors. *American Sociological Review,* 57**,** 198-213.

Qian, Nancy 2008. Missing Women and the Price of Tea in China: The Effect of

Sex-Specific Earnings on Sex Imbalance. *The Quarterly Journal of Economics*, Aug.

Rutkowski, J. 2001. Earnings inequality in transition economies of Central Europe:Trends and patterns during the 1990s. *SP discussionpaperno.0117. Washington,DC:TheWorld Bank.*

Seguino, Stephanie and Grown, Caren 2006. Gender Equity and Globalization: Macroeconomic Policy for Developing Countries. *Journal of International Development*, 18**,** 1081–1104.

Silles, Mary A. 2010. The implications of family size and birth order for test scores and behavioral development. *Economics of Education Review,* 29**,** 795-803.

Solon, Gary, Barsky, Robert and Parker, Jonathan A. 1994. Measuring the cyclicality of real wages: how important is composition bias? *NBER working paper*.

Stock, J. H. and Yogo, M. 2002. Testing for Weak Instruments in Linear IV Regression. *NBER Technical working paper 284*.

Thomas, Duncan 1990. Intra-Household Resource Allocation: An Inferential Approach. *Journal of Human Resources,* 25**,** 635–664.

Uusitalo, R. 1999. Returns to Education in Finland. *Labour Economics,* 6**,** 569-580.

Von Braun, Joachim 1988. Effects of Technological Change in Agriculture on Food Consumption and Nutrition: Rice in a West African Setting. *World Development* 16**,** 1083–1098.

Weesie, Jeroen 1999. sg121: Seemingly unrelated estimation and the cluster-adjusted sandwich estimator. *Stata Technical Bulletin. Reprinted in Stata Technical Bulletin Reprints, vol. 9, pp. 231–248. College Station, TX: Stata Press.,* 52**,** 34–47.

Zhang, Junsen, Zhao, Yaohui, Park, Albert and Song, Xiaoqing 2005. Economic returns to schooling in urban China, 1988 to 2001. *Journal of Comparative Economics,* 33**,** 730-752.