Family Size and Intergenerational Income Mobility: Evidence from China's One-Child Policy

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ABSTRACT This paper examines the effect of family size on intergenerational income mobility by exploiting the plausibly exogenous variation in fertility caused by the One-Child Policy in China. This paper finds that 2SLS estimates for the family-size effect are both statistically and economically distinguishable from the OLS estimates, indicating the endogeneity in the single equation model. The IV estimations reveal a positive but insignificant effect of family size on intergenerational income mobility, and this result tends to be unaffected by transitory shocks in earnings and life-cycle bias.

1. Introduction

A large body of studies in economics and other social science disciplines has long been interested in intergenerational mobility. Measured by the elasticity of child income with respect to parental income, the intergenerational income mobility provides an important perspective to examine the degree of equal opportunity in society. Due to the raising interest in its underlying causal mechanisms, earlier studies have discussed a number of factors that may account for the degree of intergenerational income persistence, including education (Restuccia and Urrutia, 2004; Pekkarinen *et al.*, 2009), public education spending (Mayer and Lopoo, 2008; Holter, 2015), family background (Bloome, 2007; Björklund, 2009), and neighborhood (Chetty and Hendren, 2016). Overall, the basic idea behind is that these factors could play a role in human capital formation of the child, and would thus lead to an intergenerational correlation of greater than zero given that returns to human capital investment should be expected in a well-functioning market economy.

However, the impact of family size on intergenerational income mobility has received much less attention in the literature. On one hand, as pointed out by the quantity-quality trade-off model, the presence of several siblings tends to dilute parental resources, i.e. the availability of parental resources on child decreases as the number of children increases given the resource constraints. Moreover, parental behavior may also be affected by an increase in family size due to childbearing effects and resource reallocation. Consequently, an additional child may lead to decreased human capital investment in the child. On the other hand, one could also imagine the case that the existence of siblings may be beneficial to the child's human capital formation due to economies of scale in children's education (Qian, 2006) or by decreasing the probability that both parents work outside home, especially mother's participation in the labour market (Angrist and Evans, 1996; Chun and Oh, 2002). Therefore, it is possible that there exists a relation between family size and intergenerational income mobility through the process of human capital accumulation and such family-size effects remain to be an empirical issue.

This paper advances the literature by addressing the endogeneity of fertility when estimating the effects of family size on intergenerational income mobility. Past literature typically makes use of multiple births and sex composition as instruments for family size to get rid of bias originating from omitted variables and simultaneity (see e.g. Black *et al.,* 2005; Angrist *et al.,* 2010; Black *et al.* 2010; De Haan, 2010). In this paper, I deal with potential endogeneity by taking advantage of the regional variation in the relaxation of China's One-Child Policy. This approach is appealing for two reasons. First, since this policy is restrictive, compulsory, and government-sponsored, there would be less concern about the strong assumption of parental preferences and choices when using sex composition as an instrument would raise.[1] Second, while past strategies inevitably focused on households with at least three children, this paper exploits the plausibly exogenous change in family size may help in improving our understanding of the average effect of an additional child in households with different family size.

The knowledge of causality is vital for public policy development. It is meaningful to understand whether family size causally influences the transmission of economic status between generations and the direction of the effect. These results may provide some insights for policy makers on fertility related policies, after all the public policies regarding social outcomes should result in an improvement of equality and well-being in the society. In addition to China's One-Child Policy, similar aggressive public promotion or even forced-sterilization programs can also be seen in a number of other developing countries, such as India, Mexico, Indonesia, and the Arab countries. Furthermore, allowing for family-size difference provides us an opportunity to learn more about the causal mechanisms driving the intergenerational income mobility. If there does exist a significant family-size effect, it might be of help to explain the pattern of intergenerational income

[1] When using sex composition as an instrument, the identification strategy relies on the assumption that parents who have n ($n \geq 2$) births of the same sex would decide to have another birth due to their preference for a mix-sex of children.

elasticity among economies with different fertility rates as well as its evolution in a specific country over time.

The main finding of this paper is that in contrast to the positive and statistically significant family-size effects on the intergenerational income mobility suggested by the OLS estimations, 2SLS analysis reveals that there is no evidence of a causal link between family size and intergenerational income mobility, which is measured as children's income with respect to their father's income. In particular, based on the data derived from the China Family Panel Studies (CFPS), the incomes used to calculate the elasticity have been adjusted from transitory shocks in earnings and life-cycle bias. [2]

The rest of this paper is laid out as follows. Section 2 provides a review of existing literature on the effects of family size and discusses its potential endogeneity issue as well as previous attempts at establishing the causality. Section 3 introduces the institutional background concerning China's One-Child Policy. Section 4 presents the data, sample restrictions and summary statistics. Section 5 describes the econometric methods. In Section 6, I discuss the empirical results. Finally, Section 7 concludes.

2. Literature Review

2.1 Family-size Effect

There is an extensive literature on the link between family size and child outcomes. Introduced by Becker (1960) and developed by Becker and Lewis (1973) and Becker and Tomes (1976), the quantity-quality trade-off model establishes a theoretical hypothesis for this issue. The model indicates that an increase in child quantity will increase the shadow price of child quality with a given budget constraint, suggesting an inverse association between quantity and quality. Two important implications can be derived from the model. First, family size is an input in the production function of child quality. Second, exogenous reductions in family size should increase parental investment in children, and thus improve children's outcomes and well-being through

[2] To do so, in addition to including age controls, this paper measures father's earnings by averaging father's income of three income years in a five-year period. As well, child's income is collected from the best available survey conducted when the children have been closer to their mid-age.

human capital accumulation. However, one could also imagine the case that siblings may benefit the child's human capital formation by such as sharing books and knowledge or decreasing the probability that both parents work outside home. Therefore, the family-size effect on child's outcomes remains an empirical issue. In empirical work, measures of child quality vary with respect to intelligence, health, educational attainment (e.g. school enrolment, grade retention, schooling year), and to labour market performance (e.g. earning and employment status), covering different periods of child's life. Using data sets from a variety of countries and employing different identification strategy, results of the family-size effect in empirical studies are mixed, both in signs and levels of significance.

Whether an increase in family size decreases child academic achievement is conflicted in previous studies. Distinct trade-offs between family size and educational outcomes are supported by Hanushek (1992), who investigated the relation between the number of siblings and their preschool and school performance in Indiana, the United States, as well as Ponczek and Souza (2012) , who examined the effects of family size on educational output and human capital formation for boys and girls as well as young female adults in Brazil. However, taking birth order into account, Black *et al.* (2005) find that although sibship size negatively correlated with child educational attainment in Norway, the effects reduce to zero once birth orders are included. Contrary to the findings in those studies, Qian (2006) argues that in China first-born child benefits from an additional child by an increase of school enrollment for 16 percent, and she suggests that economies of scale in schooling may be an explanation of the results.

In terms of the labour market performance of the children, Kessler (1991), Angrist *et al.* (2010), and Åslund and Grönqvist (2010) find no evidence for negative consequences of an exogenous increase of sibship size on employment status, level or growth of earnings, or welfare dependence in the United States, Israel and Sweden respectively. In contrast, Björklund (2004) observes that those who have one sibling earn 10% more than those who have four or more siblings in Sweden, Finland, and Norway. Also, a statistically significant, positive relation between the exogenous increase in family size and labor force participation for boys and girls is found in Brazil by Ponczek and Souza (2012).

The role of family size on child's IQ is examined by Black *et al.* (2010) on the basis of multiple data sets in Norway. Interestingly, they found conflicting results by using different instruments. As a conclusion, they suggested that there are likely no significant, negative effect on children's IQ of *expected* increase in family size, based on the estimates obtained when sex composition is used as an instrument; however, the *unexpected* shocks to an increase of family size resulting from twin births on the IQ scores of existing children are negative and significant.

On the other hand, a line of literature sheds light on the impact of family size on parental outcomes. In particular, these studies reflect on the effect of childbearing behavior on maternal outcomes, such as labour supply, health, and other socioeconomic consequences of married as well as unwed women. Using the instrumental variables method to derive plausibly exogenous sources of change in family size, it is argued that a larger family size leads to a reduction in female labour force participation and also a higher probability of unhealthy status, in terms of weight and blood pressure (see e.g. Bronars and Grogger, 1994; Angrist and Evans, 1996; Chun and Oh, 2002; Wu and Li, 2012). Intuitively, these effects of family size on parental outcomes might further affect their investment in their children and thus affect the child's future outcomes.

While the empirical studies of the relation between family size and the outcomes of children as well as their parents are ample, there is little previous literature which focuses on the transmission between the two generations or investigates the intergenerational income elasticity by family size. Such endeavor is highlighted by Lindahl (2008), who used a sample of nearly 74,000 individuals born in Sweden between 1962 and 1964 to examine the birth-order and family-size effects on intergenerational income mobility. He found significant birth-order and family-size patterns in the transmission of economic status between fathers and sons. Specifically, the intergenerational income elasticity tends to decrease with family size as well as birth order in a given family size, and these differences cannot be explained by differences in the age of the fathers at the time when their income data were collected. [3]

[3] He also conducted the analysis for father-daughter, mother-son, and mother-daughter samples. However, there are no significant differences in the elasticity connected with family size in these samples.

2.2 Endogeneity of Family Size

The principal difficulty along the way is to provide evidence on the causal effect of fertility on child outcomes. Potential correlation between family size and the error term is the main concern. This is because some unobservable parental characteristics, such as ability, wealth expectation, household environment and culture factors, may affect parents' fertility decisions as well as their investments in children and eventually affect their children's future outcomes. For instance, parents with relatively less endowment of ability and wealth might have limited knowledge and access to contraceptive methods resulting in having more children and meanwhile they tend to invest less in their children. In this case, ordinary least squares (OLS) estimates will be downward biased. In contrast, if a positive shock to parents' incomes, or a positive expectation on the family's future income, increases the likelihood of having another child and at the same time increases parental investment in their children, we face an opposite problem — the family-size effect will be overestimated. Another concern simultaneity. While family size could affect children's outcomes, the quality of the previous children may also shape parents' perception of the desired number of children. Overall, both omissions of explanatory variables and simultaneity lead to endogeneity, in which case OLS estimates will become biased and inconsistent and can no longer be interpreted as measuring causal effects.[4]

As a result, any attempt at disentangling the causal mechanisms linking family size and child outcomes should take potential endogeneity of fertility into consideration. One empirical strategy is to use siblings and difference out family-level fixed effect (see e.g. Guo and VanWey, 1999)[5]; however, the validity of this technique receives some doubts due to its strong assumptions about parental decisions.[6]

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[4] With regards to the studies on parental outcomes, endogeneity of fertility is also a concern. The joint determination of fertility and labour supply is the main source. See Angrist and Evans (1996) for a discussion in greater detail.

[5] Note that the single equations for each of the siblings in a sibling pair must be established for different times to allow a change in sibship size so that the family influences could be cancelled out. This analysis assumes that the time-invariant family influences shared between the two members of siblings must be same, and thus the family size effect on both of them as well as in the diffidence model are equivalent.

[6] Phillips (1999) has provided thorough comments on the believability and generality of the work by Guo and VanWey (1999) and discussed some explanations for their results.

Recently, a burgeoning empirical literature resorts to the instrumental variable (IV) strategy to address the endogeneity issue. The success of the validity of this identification approach depends on to find an instrument which is correlated to the treatment assignment but not correlated to the error term. In the context of family-size effects, we therefore need a source of variation in fertility, which is orthogonal to any unobservable characteristics of the household, to serve as the instrument. Two instrumental variables for family size widely used in the existing literature are multiple births and sex composition.

The idea to use the birth of twins as an instrument is first proposed by Rosenzweig and Wolpin (1980). They assumed that an optimal number of children is desired by parents, and thus unplanned twins birth causes an exogenous shock to the size of a household due to its randomness and perfect compliance. In practice, the instrument is usually constructed as a dichotomy variable indicating whether the *n*th child is a multiple-birth, and samples are restricted to children born before birth *n* and sometimes further to the first-born children only. Recent estimates using the birth of twins include those by Black *et al.* (2005), Cáceres-Delpiano (2006), Angrist *et al.* (2010), Ponczed and Souza (2012).

Arguing the phenomenon of parental preferences for a mixed-sex of children, sex composition is another instrument constructed in recent literature (see e.g. Angrist and Evans, 1996; Conley and Glauber, 2006; Angrist *et al.*, 2010; Black *et al.* 2010; De Haan, 2010).[7] Its main idea is that parents who have previous *n*-1 children all of the same sex are more likely to have an *n*th-child than equivalent parents with opposite sex children; as a result, if the sex composition of children is randomly given by nature, families that randomly had multiple same-sex children will be more likely to deliver another birth. Researchers thus construct a binary indicator for the *n*-1 children being the same sex as the instrument and look at the performance of the *n*-1 children. [8]

[7] There is also some literature using the sex of first birth as the instrument in response to son preference. For example, Chun and Oh (2002) uses this instrument to look at the effect of fertility on labour force participation of married women and found that an additional child reduces the labour force participation of married women in Korea by 27.5%.

[8] One concern with this identification strategy is that the sex composition may have a direct effect on children's educational attainments. In empirical studies, Conley (2000) and Deschênes (2007) find negative effects; however, Kaestner (1997) finds no effect of sibling sex composition on the educational attainment of white males or females.

A limitation of both sources of variation above is that they fail in capturing the effect on the marginal child. Specifically, the group of studies is focusing on the older children who have an extra younger sibling so that such strategies cannot identify the effect on the younger child of being born into a larger family. Another limitation is that as an implicit requirement by these instruments, samples are restricted to the families who has three children in minimum, hence the effect of increasing the number of children from one to two still remains unknown.

In recent work, China's One-Child Policy has been explored by researchers for re-visits of the quality-quantity trade-off theory. For example, based on the China Health and Nutritional Survey (CHNS), Qian (2009) makes use of the policy to estimate the effect of family size on school enrolment of the first child and Liu (2013) uses it on child height. Furthermore, on the parental side, this policy has also benn exploited to examine the causal identification of family size on maternal health outcomes by Wu and Li (2012). Amongst this arm of literature, several features related to the policy enforcement are exploited for the construction of instruments, including the eligibility for having another birth if the first child is a girl, the amount of fines for above-quota births, and the eligibility for having two children.

3. Institutional Background

With the purpose to curb the rapid growth of population after the famine during the Great Leap Forward[9], a series of birth control campaigns were commenced in China in the 1960s. Subsequently, the government's interest in family planning heightened considerably in early 1970s, with the issuance of "Later, Longer, and Fewer", where "Later" encouraged later marriage suggesting 23 years old for female and 25 years old for male, "Longer" required the spacing of three years between pregnancies, and "Fewer" implied that a couple should have no more than two children. Technically, it was still voluntary at this stage when the family planning policy was implemented. In 1979, famous known as the One-Child Policy, the Chinese government initiated the strictest national policy ever adopted. With the inclusion of family planning into the Constitution, previous persuasion has been replaced by directly targeting the number of children

[9] The Great Leap Forward was an aggressive campaign aiming at agricultural and industrial sectors launched by the Communist Party of China from 1958 to 1962.

per family. The policy was relentlessly enforced by heavily penalizing the couples for unsanctioned births. The high fine is a predominantly common form of penalty. Moreover, for employees in the public sectors, such as state-owned enterprises and governments, a violation of the policy would jeopardize their employment status, opportunities for housing, and chances of promotion. People would also lose benefits or rewards associated with having only one child, which take the form of different subsidies. Furthermore, unsanctioned children may also be discriminated in terms of education and health care.

One should note that the One-Child Policy does not necessarily mean that every household should only have one child. Actually, the fertility eligibility set by local governments is varied across communities, especially after its relaxations since 1984. Generally speaking, the birth limits tend to be more rigid in urban areas than in the suburban, and more rigid in east regions compared to the west. As well, ethnic minorities enjoyed more relaxations compared to the Han. Also, considering the sex preference for boys in China, in some communities, a second or third child is allowed if the parents have previous children all are girls. Further, some sparsely populated regions have the eligibility to have two or three or even more children regardless the sex composition of their existing offspring.

4. Data

Data used in this paper is the microdata samples derived from the 2010 to 2014 waves of the China Family Panel Studies (CFPS). The CFPS, administered by the Institute of Social Sciences of Peking University, is a longitudinal survey commenced in 2010 with a nationally representative sample of 635 communities, 14,798 families, and 33,600 individuals. All individuals over age 9 in every single sampled household constitute the core samples of the CFPS and have been received follow-up surveys on a two-year basis. With its purpose to reflect the dynamic performance of people's wellbeing in China, the CFPS is a great source for intergenerational research in a developing country context. This is because it has constructed a dataset specifying family relations and followed all core samples in these families so that it is possible to link the children's income to their parents over multiyear measures. Most importantly, the survey provides detailed community-level information on the enforcement of the One-Child Policy, which can be exploited as exogenous variation in family size.

In order to look at the impact of family size on intergenerational income mobility, I first matched each of the children to their parents in the original data through a unique identification number. The CFPS survey provides five data sets, i.e. family roster[10], child and youth[11], adult[12], family, and community, on the levels of individual, family, and community. Information on family relations are provided in the family-roster data set of the 2010 baseline survey, however detailed information for individuals of both generations are limited here. In order to obtain detailed personal information for analysis, such as income and demographic and geographic characteristics, these data should be combined from the adult database. Therefore, I build up my sample on the basis of the 2010 CFPS baseline family-roster data. Based on that, parental incomes and a set of parental characteristics from the individual survey are linked to the child with the parental identification number, and the child's income and a set of other personal characteristics are merged to the child from the individual survey with the child's personal identification number. Note that the databases on individual level are provided as cross-sections in each wave of the CFPS survey, so different variables are extracted from the different wave of individual surveys depending on needs.

Income is measured as total income. The total income defined by the CFPS comprises wages and salaries, income from agricultural production and non-agricultural business, income from property, government transfers, and other incomes. For the younger generation, income is measured in 2014, while the parental income is measured as father's average income in the years of 2010, 2012 and 2014, and all incomes are inflation-adjusted to the 2014 Chinese yuan.[13] Following the massive literature on intergenerational income elasticity, I use father's income to examine the income correlation between two generations in order to make the results of this paper more comparable to findings from previous studies. Measurement error in earnings is a vital issue along the process of obtaining precise estimates of elasticities. On father's side, if father's *permanent* incomes are measured with errors, this could lead to biased estimates. In the case where father's earning using is measured using data collected from a single year, transitory shocks may cause measurement errors. Solon (1992) and Zimmerman (1992) suggest using average earnings for 4 or 5 years for improvement. In light of this, this paper averages father's incomes in 2010, 2012, and 2014, and

[10] The family-roster data set describes the relations within a family and some basic information of the members of the family, including their personal identification numbers.
[11] The child-and-youth data set consists of the individuals who are less than 16 years old.
[12] The adult data set consists of the individuals who are 16 years old or older.
[13] The CPI data derived from National Bureau of Statistics of China is used to adjust the price.

uses the average earnings in this five-year period as the measure of father's income so as to reduce the measurement error. Moreover, if the child's income contains classical measurement error, the measurement error on the right-hand side of the equation should not bias estimates but would affect the standard errors of the OLS estimates.[14] In addition to persistent transitory variation mentioned above, lifecycle bias is another source of the measurement error. As suggested by the theoretical and empirical evidence, the lifecycle bias is smallest when incomes are measured around mid-career life (see e.g. Haider and Solon, 2006; Nybom and Stuhler, 2016). This paper measures the child's income in 2014, although the income data in the earlier surveys are also available. This is because the younger generation has been older and closer to their mid-age in 2014, compared to when they were interviewed in the previous surveys. Meanwhile, this also results in the inclusion of more well-educated individuals, avoiding the bias to those who have fewer years of schooling. The set of variables specifying personal characteristics of both generations are extracted from the surveys on the individual level includes their age, gender, education, urban status, province and ethnic group. In terms of age, offspring's age is measured in 2014 while father's age is an average among his positive income years, thereby the age variable could have a better description on the income variable during analysis.[15]

While the information on family size can be directly obtained from the original CFPS data, the exact birth order of each child needs further construction. Several steps are applied: first, detailed information of each of a father's children, including the date of birth and personal identification number, is provided in the father's individual survey, so that all of the father's children can be ordered. Then, the information for a father's children is linked to the child and this can be regarded as the information consists of himself/herself as well as all of his/her siblings. Finally, by judging whether the child's personal identification number could be matched to a specific personal identification number among those identification numbers which have already been ordered for all children in the family, the child's birth order can be identified.

The instruments used in this paper will be exploited from the enforcement and relaxation of the One-Child Policy among different regions in China. Details of the policy information can be

[14] The case of non-classical measurement error will be discussed later when introducing the empirical strategy.
[15] But father's age in summary statistics will be shown as his actual age in the year 2014.

extracted from the community survey of the 2010 wave, including the number of children eligible for a family, the number of children eligible for a family without a son, and the minimum penalty for violating the family planning policy.[16] Individuals are then matched to these policy variables by their unique community identification numbers. Considering that the fines levied on the violating family has experienced several alternations during these years, and that the considerable amount of missing values of this variable in the data set, the minimum violation penalty will not be exploited as an instrument for family size.

A number of sample restrictions are imposed on my matched data set. My sample consists of non-twin children born in the years 1976 to 1994 and their parents, who were born after 1949. To begin with, since each of these individuals is matched to their fathers through a unique parental identification number, the sample is inherently restricted to the adult individuals whose fathers have a valid personal identification number, i.e. those who have completed the individual survey. A total of 2990 father-children links are matched from the original CFPS data. The sample of this paper is restricted to individuals who are ranged from 20 to 64 years old (discarding 636, or 21.3%, observations). The lower bound of the age restriction on the younger generation starts at 20, so individuals born after 1994 are excluded. The upper age bound for parents ends at 64. More exactly, this age restriction is applied to father only, since the parental earnings in this paper will be measured as father's income. The end of 64-year-old seems reasonable because this is the age before the traditional retiring age of male employee 65, and people are less likely to remain employed beyond the retiring age and may turn to retirement smoothing work practice. Also, for the self-employers, who mostly work in the agriculture sector in this data set, they tend to cease working at this stage of life. As a result, individuals whose father born before 1949 are excluded since they had been at least 65 years old in the 2014 survey. Note that the siblings of the individuals in the sample may be born after 1994; however, one can expect this should not affect the analysis. The reason is that it was until 2015 that the One-Child Policy further relaxed the one-child restriction to all parents, when the parents in the sample tended to have completed their fertility

[16] In the 2010 community survey, three questions regarding the One-Child Policy were asked on the community level: How many births were allowed by the One-Child Policy in your community during those years? How many births were allowed by the One-Child Policy if the couple had no son in your community during those years? How much fine was imposed for violating the policy in your community?

decisions. Moreover, to account for the family planning policy, two more restrictions are imposed. One is to exclude the individuals whose oldest sibling born before 1976 (5.1% of the sample), four years prior to the implementation of the rigorous One-Child Policy, because their parents child-conceiving decision should not be affected by the One-Child Policy. Four-year is chosen because the government encouraged spacing of three years between pregnancies in that early stage of the birth planning campaign. The other is to exclude those who are multiple births (further dropping 23 observations), since they should not be affected by the family planning regime and also close birth spacing effects may confuse the comparability of the results to other individuals.[17]

With regards to incomes, this paper only includes the individuals who have a positive income in 2014, meanwhile requires the individual's father has a positive average income over the five-year income period.[18] The choice whether to require the individual's father to have a positive average income during the income years or a positive income in every single year needs careful consideration, because there exists a trade-off between better measures of father's permanent income and potential sample bias. When averaging as many years of income as possible, this average income will become a better proxy for permanent income, and therefore will lead to a more precise estimate. However, as the sample is further restricted to those have positive incomes in every single year, the reduced sample excluding those who have experienced unemployment or zero annual income may become biased towards high-income earners, since low-income workers are likely to suffer most unemployment or zero income. Such bias may in turn have dual impacts on the estimates. For one thing, as demonstrated in studies (see e.g. Österberg, 2000; Lindahl, 2008; Chen *et al.* 2017), the intergenerational income mobility for father and son tends to be lower amongst high-income fathers. For another, family size can be endogenously determined by parental income, so the estimates of family-size effect may alter if the high-income fathers are over-sampled. With the consideration of the above, the father's income is required to be positive during the observed period rather than in every single year. The income restriction reduces 32.2%

[17] An individual is identified as a multiple-birth if the individual has a sibling who was born in the same month of the same year as the individual, given that the information regarding the date of birth is only provided with the year and the month in the original data.
[18] By constructing dummies indicating if the father had positive income in a specific year, the father's number of years with positive income can be identified. Then the father's average income is calculated by the sum of his incomes during the years of 2010, 2012 and 2014 divided by his number of years with positive incomes.

of the sample, dropping 913 observations with non-positive child income and 50 observations with non-positive father's income. I also exclude individuals whose age difference between their parents and themselves is less than 17 years (3 observations), considering this case is nearly impossible to happen nowadays. Finally, this paper excludes those who live in the community where information for the family planning policy is missing due to a response of not applicable or unknown (17 observations). After imposing these restrictions, I am left with 1196 observations and 1088 father-children links.

Table 1 provides summary statistics by family size and all results are weighted. We begin with looking at the full sample. Female children account for 42.6% in my sample, implying that there are more male children earning positive incomes in the Chinese labour force. A glance at the incomes shows that children on average earn 26.7% more than their fathers with a larger standard deviation. Children are much more educated compared to their parents. While 15.5% fathers are illiterate or semi-illiterate, the figure for children drops to 1.8%. The rate of basic-education[19] completion rises to 87.4% for the children while it is 60.7% for the fathers, showing an increase of 26.7% between generations. Those who completed senior school or college constitute the largest population of the younger generation amongst all the education categories, accounting for 50.3% while most fathers have their highest educational levels at junior school, accounting for 37.5%. The proportion of those obtained a university or higher degree is much larger in the younger generation than that among the fathers, accounting for 11.8% and 0.7% respectively. In the respect of years of schooling, children on average receive 11.92 years of education, which is 3.98 years more than their fathers. The means for age are 25.45 among the children with the standard deviations of 4.56 and 51.99 among the fathers with the standard deviations of 5.94. On average, every father has 2.04 children in the sample. Next, we turn to the performance of children and fathers with different family size. We can see that the 2-children household comprises 47.1% of the sample, followed by the 1-child family, accounting for 31.3%, and then by the household has more than 3 children, accounting for 21.7%. In general, two patterns in incomes of both generations are exhibited: for one, children earn more than their fathers with a range of 14.4 percent to 67.8 percent. In particular, the most dramatic rise in incomes between generations is found in

[19] Basic education refers to the education at both primary and junior schools since these nine years of education are covered by the national compulsory education law in China enacted in 2006.

the 3-and-more-children families. Meanwhile, both generations' means of income decrease with family size. It is noted that children who have sibling(s) earn less than the average level. But the variation in children's incomes with different family size is smaller than that in fathers' income. With regards to educational attainments, several patterns highlight. First, there shows a decline in years of schooling for both fathers and children when family size increases. Second, in each category of the lower education levels, which are junior school or less, the share of children increases with family size. In contrast, the proportion of children decreases with family size in each category of the higher education levels, which are senior school or college or more. Third, the distribution of fathers' education levels displays a very similar pattern. The latter two turn out to be consistent with the identification concern that parents with more education may prefer a smaller family size and value investment in their children more. Another noteworthy point is that there is much less female in the sample of 1-child families while the sex ratio remains very consistent among the full sample and other sub-samples. This sex ratio imbalance might imply some sex selection in the 1-child families,[20] which is consistent with previous findings (see e.g. Attane, 2002; Li, 2011). Furthermore, for both generations, the statistics for age and ethnic group remain rather consistent amongst different family size.

5. Empirical Strategy

This paper estimates the effect of family size on intergenerational income elasticity. To deal with the issue of endogeneity of family size, the identification strategy relies on an instrumental variables approach and exploits the variation of the fertility eligibility imposed by China's One-Child Policy across localities.

I start with the standard model of intergenerational mobility. The focus of intergeneration economic status persistence lies in the correlation between the incomes of the two generations.[21]

[20] One may argue that the imbalance sex ratio could arise from sample restrictions when excluding those children who had non-positive income in 2014. However, this should not be the case. First, if this is true, it implies that female children from one-child families are more likely to have non-positive income compared to those female children from families with a larger size, and this does not make sense. Moreover, among those 913 observations dropped, only 351 are female (i.e. 72 from families with one child, 181 from families with two children, and 98 from families with three children and more), while 562 are male (i.e. 119 from families with one child, 270 from families with two children, and 173 from families with three children and more).

[21] Note that what interests us here is the relation of correlation rather than the casual link.

Most research estimates this association as follows

$$lny_c = \beta_0 + \beta_1 lny_p + \mu_i \quad (1)$$

where the subscript c represents the young generation, the subscript p represents the parental generation. lny_c is the natural logarithm of the child generation's income, and lny_p is the natural logarithm of parental income. Specifically, the prototypical approach in the literature is to regress son's income on that of father's. [22] β_1 is the intergenerational income elasticity, which measures the percentage change in the offspring's income with respect to a marginal percentage in father's income, and thus $(1 - \beta_1)$ is a measure of the intergenerational income mobility. β_1 falls in the range between 0 and 1. When β_1 equals to 0 (and mobility equals to 1), it implies that father's income has no effects on a child's economic success, so the economic status in the society is fully mobile; and vice visa.

The major challenge in this line of research is that the earnings should be a measure of permanent incomes. As already discussed in the data section, the measurement error, either classical or non-classical, in the explanatory variable (i.e. the proxy for father's permanent income) will cause biased and inconsistent estimates. While measurement error in the dependent variable (i.e. the proxy for the offspring's permanent income) should remain unbiased and consistent estimates but result in larger standard deviations if the measurement error is classical. In the case of non-classical measurement error, in which the error is correlated with the true variables of interest or correlated with the measurement errors in the explanatory variables, mismeasurement in the dependent variable may cause biased estimates; however, this issue could be fixed by employing the instrumental variables approach, which will be introduced later, if these measurement errors are not correlated with the instruments. In order to reduce the measurement errors that arise from transitory shocks in earnings as well as life-cycle bias, this paper uses the average income over the income years as the proxy for father's income, and uses income in 2014, when more individuals

[22] This is because the whether the difference of the intergenerational income mobility between father-son and father-daughter is significant is argued by studies (see e.g. Österberg, 2000). This paper does not distinguish between son and daughter in the younger generation considering that the goal of this paper is to investigate the relation between family size and intergenerational income mobility rather than to obtain a precise estimate of the intergenerational income mobility itself. But sex will be added as a control variable in the preferred model.

in the younger generation have reached their mid-age, to measure the younger generation's earnings. In addition, on account of the large standard deviations shown in the age of both generations, I also include age controls for both father and child in the equation to adjust for the life-cycle variation in earnings, as suggested by Solon (1992). The equation this paper uses to estimate intergenerational income elasticity is written as

$$lny_{ci} = \beta_0 + \beta_1 lny_{pi} + \gamma A_i + \mu_i \quad (2)$$

where lny_{ci} is the natural logarithm of the individual i's income in 2014 and lny_{pi} is the natural logarithm of the individual i's father's average income in the years 2010 to 2014, and A_i is a vector of age controls to adjust for the life cycle variation of both generations, which includes the father's average age during his positive income years, the squared of father's average age over his positive income years, the individual's age in 2014, and the squared of the individual's age in 2014.

The objective of this paper is to examine the role that family size plays in intergenerational income mobility. To determine whether family size affects the elasticity of the individual's earnings with respect to parental earnings, specifically father's earnings, I add family size and its interaction term with father's income in the model based on Equation (3). The main equation of this paper takes the following form

$$lny_{ci} = \beta_0 + \beta_1 lny_{pi} + \beta_2 size_i + \beta_3 (size_i \cdot lny_{pi}) + \gamma A_i + \varphi X_i + \mu_i \quad (3)$$

where lny_{ci} denotes the natural logarithm of the individual i's income in 2014, lny_{pi} denotes the natural logarithm of the individual i's father's average income in the years of 2010, 2012 and 2014. $size_i$ is the family size of the individual i, which is calculated as the number of his or her siblings (includes the deceased) plus the number of one, $(size_i \cdot lny_{pi})$ is an interaction term between family size and father's average income during his income years. A_i is a vector of age controls for both generations as in Equation (2), X_i is a vector of controls for the characteristics of the individual i, including birth order, gender, ethnic group (2 categories), urban status (2 categories) and province-fixed effects (25 categories). The intergenerational income elasticity is $\beta_1 + \beta_3 \cdot size_i$, and the intergenerational income mobility is $1 - (\beta_1 + \beta_3 \cdot size_i)$. The parameter of interest is β_3, which measures the family-size effects on intergenerational income mobility. If family size decreases intergenerational mobility, β_3 will be positive; and if family size increases intergenerational mobility, β_3 will be negative.

As mentioned earlier, I am likely to face an endogeneity issue when estimating Equation (3) by OLS, since fertility, or say the number of children, and the investment in children can be jointly determined by parental preferences. OLS estimates could be downward biased in the case where parents who value human capital investment also prefer smaller family size, while will over-estimate the family-size effects if larger households benefit children's human capital accumulation. To address this issue, I exploit the plausibly exogenous variation in the number of children caused by the strict enforcement of the family planning policy in China, which is also commonly known as the One-Child Policy. Due to its rigorous implementation with harsh punishments as I mentioned earlier, it is believed that parents' fertility decision was exogenously determined by the eligibility under the regime. The feature of the policy that the fertility eligibility for a household differs among localities provides us a favorable quasi-natural experiment to look at the family-size effect.

I exploit two variables derived from the community survey to construct instruments for the family size. First, the variable specifies the number of children that parents are eligible to have can be directly identified as an instrument, and this instrument is applicable to all individuals in the sample, whatever their sex and birth orders. Another is the variable related to the relaxation for one more child if the couples' first or both of their first and second children is (are) girl(s). The eligibility for having 1-son-2-child and 1-son-3-child is jointly determined by sex and region. Parents are considered as eligible for having one more child should meet the conditions that the family planning policy enforced in their region allow the relaxation that if their previous child (children) is/are (both) a girl, they were eligible for one more child. Only the combination of these is exogenous to those existing children. Consequently, this identification strategy is applied to the first-born child in all regions and to the first-born and second-born children for those regions with 1-son-3-child eligibility. Hence, the instrument should take birth order, sex, and the extent of relaxation into account. I therefore construct the instrument as an interaction term of the parents' previous children, which accounts for both birth order and sex, and the family size eligibility for no-son parents. The first stage equations can be written as

$$size_i = \beta_0 + \beta_1 policy_i + \beta_2 (girl_i \cdot policyson_i) + \beta_3 (policy_i \cdot lny_{pi}) + \beta_4 (girl_i \cdot$$

$$policyson_i) \cdot lny_{pi} + \gamma A_i + \varphi X_i + \mu_i \tag{4}$$

$$size_i \cdot lny_{pi} = \beta_0 + \beta_1 policy_i + \beta_2(girl_i \cdot\ policyson_i) + \beta_3(policy_i \cdot lny_{pi}) + \beta_4(girl_i \cdot$$
$$policyson_i) \cdot lny_{pi} + \gamma A_i + \varphi X_i + \mu_i \tag{5}$$

where $policy_i$ is the eligible number of children of a household where the individual i born, $girl_i$ is a variable indicating if the family's previous births are female, i.e. the first-born child is a girl or both the first-born and second-born children are girls, $policyson_i$ is the eligible number of children if the parents do not have a son, lny_{pi} is the natural logarithm of the individual i's father's average income in the years 2010, 2012 and 2014. A_i and X_i are vectors of control variables as defined in Equations (2) and (3).

With regards to the validity of the instruments, the IV estimator's good asymptotic properties rely on two assumptions: one is that the original explanatory variables should be highly correlated with the instruments, and the other one is that instruments should be uncorrelated with the error term. For the former, due to the heavy fines levied on unsanctioned births and other forms of punishments on parents for a violation, it is reasonable to believe that families tended to comply with the policy and thus the eligibility allowed by policy should be strongly correlated with family size. Also, the impact of the One-Child Policy on fertility has been confirmed by studies, arguing that the policy has a significant effect in decreasing the fertility rate after controlling for socioeconomic development (see e.g. Li *et al*., 2005). For the latter, while the unobservable variables are on the individual level, it is believed that the birth quotas imposed on the community level required by a government-sponsored regime should not be correlated with those.

My models are built on equation (3) and will be estimated by adding the vector of controls sequentially. In the first specification, no controls for the child's personal characteristics are added, so the log of child's of income is now a function of the log of father's income, family size, the interaction term between the log of father's income and family size, and the age controls for life-cycle bias adjustments. In the next specifications, I control some characteristics of children that may account for differences in their income and that might be correlated with family size. Sex and birth order are added as controls in my second specification. While the reason for adding the sex control is mentioned earlier, the reason for including birth order is that a number of previous literature has discussed the effects of birth order on a number of children's outcomes, such as educational attainment, wages, labour market participation, and intergenerational income mobility as well (see e.g. Kessler, 1991; Black *et al.*, 2005, Lindahl, 2008). Ejrnæs and Pörtner (2004) conclude that birth order may affect children's economic status mainly through three channels (i.e. constraints, household environment, and cultural factors). Considering the potential birth-order effects on children's outcomes as well as the close relevance between birth order and family size, the variable of birth order is thus added in the specification. The third specification additionally includes urban status and ethnicity, with the consideration of their various socioeconomic features and culture factors. Finally, in light of the considerable difference in economic development among provinces, province-fixed effects are added in the fourth specification.

6. Results

Table 2 reports the regression results for the four specifications of the main equation using the estimation methods of OLS and 2SLS respectively. When estimating by OLS, the coefficient estimates for the interaction between the family size and the logarithm of father's income remain quite consistent in terms of both magnitude and signs amongst the four specifications, which implies the stability of the model. All the four estimates are negative, statistically significant at 10% level or above and of some economic significance with a magnitude ranging from -4.0% to -6.1%. By comparing the results for the four specifications, one can see an increase in the R-squared when more control variables are included. Hence, specification (4) has more explanatory power and more flexibility due to its inclusion of the full set of controls (i.e. the child's sex, birth order, urban status, ethnicity, and province-fixed effects). Therefore, specification (4) will be used as the preferred model for this paper's analysis. Results show that after controlling for the younger generation's sex, birth order, urban status, ethnicity, and province, an additional child in the family decreases intergenerational income elasticity by 4%, and this effect is statistically significant at a 10% level. In other words, the OLS results suggest a higher intergenerational income mobility with a larger family size. This inverse relation between family size and intergenerational income elasticity, i.e. the positive relation between family size and intergenerational income elasticity, is consistent with the previous finding by Lindahl (2008), who indicated that the intergenerational income elasticity tends to decrease with family size based on the labour-income analysis of Swedish fathers and sons.

However, once the potential endogeneity of family size is taken into account, the family-size effects on intergenerational income mobility are no longer significant. In my preferred model, although the sign of the 2SLS estimate remains consistent as the OLS estimates, the coefficient estimate for the interaction term between the family size and the logarithm of father's income is -0.011, implying that such effect is small and with little economic significance. Also, in terms of statistical significance, the 2SLS estimate for the interaction term does not exhibit a significance level of 10%. Further, a closer look at the results finds that the standard errors of the estimate are rather high (0.118). This implies that if one would like to reject the null hypothesis at the significance level of 5%, the coefficient estimate should be around -0.059, which will be 5 times bigger than the actual case. Regression results for first stage equations are presented in Table 3A and 3B. One important assumption underlying the reliability of IV estimation is that the instruments should be highly correlated with the endogenous explanatory variables, otherwise there would be serious bias in estimates. To test this, applying Staiger and Stock (1997)'s rule of thumb, the instruments for both family size and the interaction term between the family size and the logarithm of father's income pass the week instrument tests because their F statistics are far more than 10, which are 27.04 and 24.35 respectively. Therefore, the instruments used in the paper are sufficiently correlated with the endogenous explanatory variables in the equations.

7. Conclusion

This paper examines the causal effect of family size on the intergenerational income mobility. Due to the joint determination between parental decisions on family size and investment in children as well as unobservable explanatory variables, the main identification problem in measuring such an effect is the potential endogeneity of family size. To fix this problem, I exploit the plausibly exogenous variation in the fertility eligibility across communities caused by China's One-Child Policy using the instrumental variables estimation approach. Due to the compulsory and rigorous birth limits enforced by the policy, it is believed that fertility eligibility imposed by the policy observed on the community level should be highly correlated to family size and uncorrelated to parental preferences and parent's choice in family-size planning.

A simple OLS analysis shows a significantly positive effect of family size on intergenerational income mobility, which is consistent with the previous study by Lindahl (2008). In contrast, when the issue of endogeneity is taken into account, the IV estimators reveal that the OLS coefficients are upward biased and the family-size effects are not statistically significant. Such distinguishable results between OLS and 2SLS estimates suggest the underlying endogeneity of family size in the single equation model. This paper finds that family size is not a causal mechanism underlying the transmission of economic status between generations. Moreover, this paper can also be regarded as an empirical examination on the extension of the quality-quantity trade-off hypothesis, and its results are similar to other recent studies (e.g. Angrist, 2010; De Haan, 2010) indicating that no evidence of a significant causal interpretation of family size on children's outcomes.

In particular, the intergenerational income elasticity in this paper is measured as the correlation of children's earnings with respect to father's earnings. Taking advantage of the panel feature of the data set I use, in addition to including age controls to adjust life-cycle bias, this paper also averages father's incomes of three years in a five-year period to reduce the measurement error of father's permanent income. Therefore, the results of this paper are less likely to be biased by the year when the father's income is collected. Furthermore, in addition to adding child's age controls, the instrumental variable estimation method I adopt may also fix the issue of measurement error in child's income, because it is reasonable to believe that the birth quotas imposed by the governments should not be correlated to the child's income in the future.

The exploration of the causal mechanisms underlie intergenerational mobility is motivated by the goal of achieving equality. In addition to efficiency, a society favors equal opportunity as social values — the idea that people who work hard and abide by rules should be fairly rewarded, regardless of family background. However, a fully mobile society with zero intergenerational correlation is not necessarily the optimum, given the fact that returns to education should act in a wholesome market economy. In order to determine a socially optimal level of intergenerational mobility, it is meaningful to understand the determinants which drive the correlation between generations. This paper amplifies this arm of studies with the evidence of no statistically significant causal relation between family size and intergenerational income mobility. This result suggests that a negligible effect and limited role for fertility policy to equalize opportunity in the society.

One caveat of the identification strategy employed in this paper is that the various fertility eligibility may not be randomly assigned to the regions. The allocation of birth limits tends to be affected by regional differences in population distribution and economic development, which might in turn to be correlated with ethnic groups and geographical factors, hence there still remain some concerns about whether the instruments are indeed orthogonal to the unobserved variables that may affect children and family characteristics. Also, as the imbalance sex ratio in the one-child families is shown in summary statistics, the potential use of sex-selective abortion may invalidate the assumption that the eligibility for the non-son families is exogenous.

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Table 1. Summary Statistics (Means and Std. dev.)

	Family size			
Variable	All	1 Child	2 Children	3+ Children
A. Children				
Income (¥)	23258.64	26298.85	22055.67	22503.69
	(19659.1)	(23140.9)	(18763.3)	(16891.5)
Illiterate/Semi-illiterate	0.018	0.010	0.014	0.034
	(0.1319)	(0.0994)	(0.1171)	(0.1827)
Primary school	0.108	0.064	0.1003	0.172
	(0.3100)	(0.2448)	(0.3007)	(0.3778)
Junior school	0.254	0.196	0.243	0.342
	(0.4355)	(0.3975)	(0.4293)	(0.4753)
Senior school or college	0.503	0.558	0.524	0.395
	(0.5002)	(0.4973)	(0.4998)	(0.4897)
University or more	0.118	0.172	0.118	0.057
	(0.3227)	(0.3782)	(0.3233)	(0.2324)
Years of schooling	11.92	12.89	11.96	10.73
	(3.4329)	(3.1680)	(3.2718)	(3.7034)
Female	0.426	0.314	0.478	0.437
	(0.4947)	(0.4646)	(0.5000)	(0.4969)
Urban	0.480	0.665	0.435	0.373
	(0.4998)	(0.4725)	(0.4961)	(0.4844)
Han	0.925	0.942	0.915	0.928
	(0.2640)	(0.2346)	(0.2797)	(0.2595)
Birth order	1.40	1.00	1.39	2.03
	(0.6443)	(0.000)	(0.4964)	(0.9229)
Age (in 2014)	25.45	25.63	25.29	25.60
	(4.5553)	(4.7217)	(4.4156)	(4.6676)
B. Fathers				
Income (¥)	18353.2	21021.0	19283.9	13407.5
	(16065.4)	(16592.9)	(17042.3)	(11693.6)
Illiterate/Semi-illiterate	0.155	0.113	0.143	0.227
	(0.3624)	(0.3172)	(0.3508)	(0.4200)
Primary school	0.238	0.204	0.261	0.225
	(0.4259)	(0.4033)	(0.4395)	(0.4186)
Junior school	0.375	0.404	0.417	0.254
	(0.4844)	(0.4913)	(0.4935)	(0.4361)
Senior school or college	0.224	0.263	0.173	0.293
	(0.4173)	(0.4410)	(0.3781)	(0.4572)
University or more	0.007	0.016	0.006	0.000
	(0.0848)	(0.1261)	(0.0777)	(0.0000)
Years of schooling	7.97	8.51	8.02	7.26
	(3.8235)	(3.6769)	(3.5639)	(4.3877)
Urban	0.387	0.590	0.344	0.258
	(0.4874)	(0.4924)	(0.4754)	(0.4384)

Han	0.937	0.952	0.929	0.9367
	(0.2435)	(0.2145)	(0.2570)	(0.2440)
Age (in 2014)	51.99	51.67	51.34	53.74
	(5.9395)	(5.9527)	(5.9074)	(5.6630)
C. Family Size				
Family size	2.04	1.00	2.00	3.28
	(0.8471)	(0.0000)	(0.0000)	(0.5698)
Observations	1196	374	563	259

Notes: Means are weighted and standard deviations are in parentheses. The sample consists of children born in the years 1976-94 and their fathers, who were born after the year 1949. For a description of the variables, see Table A1.

Table 2. Regression Results: Effects of family size on intergenerational income mobility

	(1)		(2)		(3)		(4)	
	OLS	2SLS	OLS	2SLS	OLS	2SLS	OLS	2SLS
Log of father's income	0.208***	0.406*	0.222***	0.303	0.205***	0.274	0.149**	0.082
	(0.076)	(0.243)	(0.075)	(0.241)	(0.075)	(0.243)	(0.077)	(0.271)
Family size	0.600**	1.369	0.665**	0.916	0.652**	0.934	0.469*	0.180
	(0.283)	(0.973)	(0.284)	(0.979)	(0.284)	(0.975)	(0.277)	(1.057)
Family size × Log of	-0.062**	-0.152	-0.063**	-0.103	-0.061**	-0.091	-0.040*	-0.011
father's income	(0.030)	(0.107)	(0.030)	(0.105)	(0.030)	(0.106)	(0.024)	(0.118)
Child's age	1.121***	1.131***	1.114***	1.117***	1.109***	1.113***	1.085***	1.081***
	(0.088)	(0.089)	(0.087)	(0.088)	(0.087)	(0.087)	(0.087)	(0.087)
Squared of child's age	-0.018***	-0.019***	-0.018***	-0.019***	-0.018***	-0.019***	-0.018***	-0.018***
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
Father's age	-0.180*	-0.196*	-0.169*	-0.183*	-0.170*	-0.178*	-0.166*	-0.161***
	(0.095)	(0.101)	(0.095)	(0.099)	(0.095)	(0.098)	(0.095)	(0.097)
Squared of father's age	0.002*	0.002**	0.002*	0.002*	0.002*	0.002*	0.002*	0.002*
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Female	-	-	-0.370***	-0.341***	-0.364***	-0.364***	-0.371***	-0.367***
			(0.069)	(0.079)	(0.069)	(0.079)	(0.069)	(0.078)
Birth order	-	-	-0.090	-0.010	-0.087	-0.092	-0.069	-0.050
			(0.065)	(0.132)	(0.065)	(0.134)	(0.064)	(0.122)
Urban	-	-	-	-	0.087	0.085	0.058	0.057
					(0.069)	(0.073)	(0.072)	(0.075)
Han	-	-	-	-	0.215*	0.213*	0.203	0.203
					(0.126)	(0.126)	(0.157)	(0.156)
Constant	-4.317*	-5.802**	-4.349*	-4.712*	-4.349*	-4.857*	-3.358	-2.802
	(2.234)	(2.775)	(2.232)	(2.793)	(2.240)	(2.792)	(2.301)	(3.020)
Province controls (24)	NO	NO	NO	NO	NO	NO	YES	YES
R-squared	0.254	0.247	0.272	0.267	0.275	0.274	0.314	0.313
Observations	1,196	1,196	1,196	1,196	1,196	1,196	1,196	1,196

Notes: The dependent variable is the natural logarithms of children's incomes. Standard errors are in brackets. * denotes statistical significance at 10% level. ** denotes statistical significance at 5% level. *** denotes statistical significance at 1% level (all two-tailed tests). For a description of the variables, see Table A1.

Table 3A. Regression Results: First stage on family size

	(1)	(2)	(3)	(4)
Policy	-0.0099***	1.513***	1.582***	1.566***
	(0.058)	(0.463)	(0.472)	(0.465)
Girl × Policy-son	-0.479**	-0.002	0.016	-0.241
	(0.237)	(0.202)	(0.203)	(0.197)
Policy × Log of father's income	-0.099***	-0.112**	-0.117**	-0.114**
	(0.058)	(0.049)	(0.050)	(0.049)
Girl × Policy-son × Log of father's	0.0544**	0.014	0.013	0.040**
income	(0.025)	(0.021)	(0.021)	(0.020)
Log of father's income	0.47	0.123	0.128	0.138*
	(0.093)	(0.079)	(0.080)	(0.079)
Child's age	-0.063	-0.031	-0.033	-0.071
	(0.060)	(0.051)	(0.051)	(0.049)
Squared of child's age	0.001	0.001	0.001	0.002*
	(0.001)	(0.001)	(0.001)	(0.001)
Father's age	0.180***	-0.010	-0.011	-0.003
	(0.065)	(0.055)	(0.055)	(0.054)
Squared of father's age	-0.001**	0.000	0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)
Female	-	0.074	0.074	0.045
		(0.070)	(0.070)	(0.068)
Birth order	-	0.722***	0.724***	0.649***
		(0.034)	(0.034)	(0.033)
Urban	-	-	0.033	0.038
			(0.044)	(0.044)
Han	-	-	0.081	0.172*
			(0.076)	(0.090)
Constant	-3.443**	-0.509	-0.657	-0.911
	(1.672)	(1.416)	(1.435)	(1.416)
Province controls (24)	NO	NO	NO	YES
R-squared	0.146	0.397	0.398	0.470
F statistics of the instruments	30.64	27.46	25.13	27.04
Observations	1,196	1,196	1,196	1,196

Notes: The dependent variable is family size. The estimation method is ordinary least squares (OLS). The omitted group may consist of children who are male, not Hans, and live in the rural area in Beijing. Standard errors are in brackets. * denotes statistical significance at 10% level. ** denotes statistical significance at 5% level. *** denotes statistical significance at 1% level (all two-tailed tests). For a description of the variables, see Table A1.

Table 3B. Regression Results: First stage on family size × log of father's income

	(1)	(2)	(3)	(4)
Policy	8.905*	8.507*	9.007**	8.648**
	(5.149)	(4.389)	(4.470)	(4.374)
Girl × Policy-son	-2.110	2.261	2.402	-0.385
	(2.224)	(1.917)	(1.924)	(1.850)
Policy × Log of father's income	-0.325	-0.450	-0.481	-0.436
	(0.545)	(0.465)	(0.474)	(0.464)
Girl × Policy-son × Log of father's	0.265	-0.102	-0.115	0.185
income	(0.234)	(0.200)	(0.201)	(0.193)
Log of father's income	1.903**	2.604***	2.634***	2.683***
	(0.872)	(0.745)	(0.761)	(0.744)
Child's age	-0.422	-0.127	-0.146	-0.528
	(0.872)	(0.480)	(0.481)	(0.462)
Squared of child's age	-0.002	0.006	0.006	0.012
	(0.010)	(0.009)	(0.009)	(0.008)
Father's age	1.408**	-0.327	-0.329	-0.200
	(0.610)	(0.526)	(0.526)	(0.506)
Squared of father's age	-0.011*	0.003	0.003	0.002
	(0.006)	(0.005)	(0.005)	(0.005)
Female	-	0.634	0.635	0.383
		(0.661)	(0.662)	(0.637)
Birth order	-	6.593***	6.618***	5.869***
		(0.318)	(0.322)	(0.312)
Urban	-	-	0.280	0.326
			(0.418)	(0.416)
Han	-	-	0.606	1.514
			(0.722)	(0.845)
Constant	-41.723***	-14.964	-16.001	-19. 234
	(15.681)	(13.416)	(13.601)	(13.324)
Province controls (24)	NO	NO	NO	YES
R-squared	0.134	0.375	0.376	0.459
F statistics of the instruments	29.03	26.40	23.61	24.35
Observations	1,196	1,196	1,196	1,196

Notes: The dependent variable is family size × log of father's income. The estimation method is ordinary least squares (OLS). The omitted group may consist of children who are male, not Hans, and live in the rural area in Beijing. Standard errors are in brackets. * denotes statistical significance at 10% level. ** denotes statistical significance at 5% level. *** denotes statistical significance at 1% level (all two-tailed tests). For a description of the variables, see Table A1.

Appendix

Table A1. Variable Definitions

Variable	Definition
A. Income	
Child's income	Income in 2014
Father's income	Average income in 2010, 2012 and 2014 and adjusted to 2014 price
B. Family size	
Family size	Number of children in the family = number of siblings + 1
C. Age	
Child's age	Age in 2014
Father's age	Average age during his positive income years of 2010, 2012 and 2014
D. Child's birth order and sex	
Birth order	Order of birth in the family
Female	Indicator variable = 1 if female; 0 if male.
E. Other personal characteristics	
Urban	Indicator variable = 1 if lives in urban area; 0 if lives in rural area.
Han	Indicator variable = 1 if ethnicity is Han; 0 otherwise.
Education	Indicator variable = 1 for highest completed level of education; 0 otherwise. 5 levels: illiterate/semi-illiterate, primary school, junior school, senior school or college, and university or more. Illiterate/semi-illiterate refers to individuals who do not have education or not complete primary school. Primary school and junior school refer to those who have obtained education at primary school or junior school respectively. Senior school or college refers to those who are senior school graduate or have a post-secondary diploma or certificate but below university degrees. University or more refers to those with a bachelor's degree or higher.
F. Policy	
Policy	Eligible number of children of a family
Policy-son	Eligible number of children of a no-son family
Girl	Indicator variable =1 if the family's previous births are girls; 0 otherwise.