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Under the one child policy regime in China: did having younger sibling(s) increase the risk of overweight and underweight status?

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ABSTRACT
The implications of having any younger sibling(s) on child overweight and underweight status under China's One Child Policy Regime are complicated by multiple factors, including potential resource dilution, the stage of economic development, changing child-rearing norm, mandated birth interval and parental son preference. Using the instrumental variable method and data from China Health and Nutrition Survey 1991–2006, we find that having younger sibling(s) generally does not affect a firstborn child's risk of being overweight or obese, neither does it increase the risk of being underweight. The findings on underweight status suggest that the favourable effect of economic growth and child rearing practice have outperformed the resource dilution effect in basic nutrition needs through the years of the study. It implies that further relaxation of the One Child Policy should not increase the nutritional risk for children.

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Child overweight; child underweight; one child policy; China; causal inference

Introduction

The One Child Policy that significantly reduced the fertility level is thought to be a leading cause of child obesity in China (Ni, 2000; Taylor, 2004). Studies on fertility and child nutrition status have established that large family size leads to child malnutrition (Balderamaguzman, 1978; Rao & Gopalan, 1969), and falling fertility significantly contributes to improved nutrition intake (Hatton & Martin, 2010). These studies mainly focused on comparing the impact of having multiple children as opposed to having one or two. We know very little about the effect of increasing the number of children from one to two or three. In China, fertility had decreased to 2.9 children per family in the late 1970s before the One Child Policy took effect (Hesketh, Li, & Zhu, 2005) and continued to decline to 1.55 in 2011 (UN Population Division, 2015). As many families throughout Asia, and particularly China, began having fewer children (Jones, 2007), the opportunity arose to compare the impact on child nutrition of having an only child to having two or three. One prominent theory is that having multiple children affects child nutrition status as a result of elevated competition for household resources. Reducing the number of siblings reduces competition for those resources (Becker & Lewis, 1973; Liu, 2014). Further, studies have

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documented that having more family resources is known to contribute to child obesity in China (Dearth-Wesley, Wang, & Popkin, 2008; Hsu et al., 2011; Wang, Monterio, & Popkin, 2002), whereas a lower level of household income strongly predicts undernutrition for Chinese children (Ge, Zhai, & Yan, 1999).

Previous studies have found that compared to children with siblings, those without any siblings are more likely to be overweight or have higher height for age (Bredenkamp, 2008; Hesketh, Qi, & Tomkins, 2003; Yang, 2007). Consistent with these findings, research has documented that children with no siblings tend to consume a higher percentage of animal foods, but a lower proportion of vegetables and fruits compared to children with siblings (Ng, 2005). On the other hand, having multiple siblings is related to undernutrition in rural China (Zhang et al., 2011). These literatures suggest that being the only child as opposed to having younger siblings has considerable impact on child nutrition status. However, it is difficult to establish the causal relationship, due to the lack of control on the unobserved family and community level heterogeneities that play a role in parental decision on sibsize(the number of children) and at the same time affect child nutrition status.

For example, parents who chose to have two children or more, authorised or not, might have more sources of untraced income, and more informal support from the family planning officials and extended family. A greater threat to the validity is that these unobserved factors could change over time. For example, families might decide to have another child when their general conditions improve. Or, if they experience a downturn in financial wellbeing, a couple might decide to have another child to ensure elder care, a reflection of the absence of a pension system and the cultural norm that children continue to serve as the primary caregivers for their aged parents and even grandparents despite recent rapid socioeconomic changes and urbanisation (Chow & Zhao, 1996; Meulenberg, 2004). Lack of control on these family level unobserved characteristics challenges the validity of directly using sibsize to identify the impact on child nutrition and weight status.

Some studies used the community-level, policy-permitted number of children per couple as the instrument variable to identify the impact (Qian, 2009; Rosenzweig & Zhang, 2009), but the policy-permitted number of children per couple (birth quota) is also linked to child nutrition status through multiple channels. The quota is closely related to local economic development and population density, and could also be a reflection of the local political and cultural environment. Chongqing, Sichuan, Jiangsu, Beijing, Shanghai and Tianjin are among the most densely populated regions and subject to the most stringent policy enforcement. Also subject to stringent policy enforcement are the richest and most developed regions or metropolitan areas, while the less developed regions are extended some leniency (Gu, Wang, Guo, & Zhang, 2007). Therefore, the exclusion restrictions criteria may not hold for birth quota to be a valid instrument.

Using the China Health and Nutrition Survey (CHNS) data collected in 1991, 1993, 1997 and 2000, 2004 and 2006, we examine the amount in monetary fines levied for an extra child across time and location as the instrument to identify whether having younger siblings affects a child's underweight and overweight status under the One Child Policy environment. Extensive analysis on whether the variation in fines is a valid instrument is conducted in discussion on methods below.

Conceptual framework

It is well documented that an increase in access to resources contributes to diminishing child undernutrition (e.g. Svedberg, 2006). Less intuitively, access to resources is positively related to child obesity in China (Dearth-Wesley et al., 2008; Hsu et al., 2011; Wang et al., 2002). One major factor that may account for the positive resource-overweight link could be the affordability of calorie-dense foods and labour saving devices. Energy-dense foods continue to have higher relative prices compared to energy light foods (Ge et al., 1999; Lu & Goldman, 2010). Empirically, higher income groups consume more snacks, and the income gap in consumption of snacks and fried foods during 1991–2004 increased (Wang, Zhai, Du, & Popkin, 2008; Wang, Zhai, Zhang, & Popkin, 2012). One study found that, as the income inequality increased in China through 1991 to 2006, the positive SES gradients of child overweight status also increased significantly (He, James, Merli, & Zheng, 2014).

According to the resource dilution model, a decrease in sibsize reduces resource competition (Becker & Lewis, 1973; Becker & Tomes, 1976; Blake, 1981; Steelman, Powell, Werum, & Carter, 2002). In the context of China, if the resource dilution effect exists, children with fewer siblings are expected to receive more resources and in turn are less likely to become underweight and more likely to become overweight.

However, as many other social/economic forces, resource dilution effect does not happen in a vacuum. Previous studies have identified some important contextual factors including the stage of economic development, total number of children per family, birth interval, whether the birth is planned and culturally acceptable child-raising norms, etc. It was suggested that development stage could modify the resource dilution effect and in turn lead to varying quantity-quality trade-off effect. Li, Zhang, and Zhu (2008) posited that well-functioning public education/health systems in developed countries absorb in large part the cost of child quality and therefore a quantity-quality trade-off effect is unlikely to be observed; whereas developing countries are more likely to witness such trade-off. Using a one per cent sample from the 2000 China census data, they found a negative effect of family size on child education. The effect might be more pronounced when fertility level is high. Bisai, Mahalanabis, Sen, and Bose (2014) found that the rate of becoming underweight was significantly higher among children who had three siblings or even more. A planned increase in a family's number of children might have less resource dilution effect than an unplanned one (Liu, 2015). And the level of parents' altruism varied across different types of families and seemed to depend on culturally acceptable practices (Desai, 1992).

Under China's One Child Policy, there are reasons to suspect that having siblings might affect the allocation of resources in a different way. On the one hand, having an only child changes the dynamics of decision making within the household, which is evidenced by findings that in Beijing those with only child status determine as much as 70 per cent of a family's overall spending compared to 40 per cent in the United States (McNeal & Wu, 1995; Ng, 2005). In such cases, having no siblings could give a child more access to resources than the resource dilution hypothesis alone would predict, thus amplifying the resource dilution effect.

However, on the other hand, as mentioned above, other factors that might have dampened the quantity-quality trade-off effect or resource dilution effect include the following.

(1) The stage of economic development in a country to a large extent determines the proportion of expenditures for food consumption which in turn could modify the resource dilution effect. If expenditures for food consumption only take a manageable portion of the family's budget, the dilution effect might be only pronounced in consumption of luxury goods, not in basic nutrition intake. Thus, having one or two more children might not affect the firstborn's nutrition intake in a significant way. The proportion of family income spent on food decreased from 57.5 per cent in 1978 to 37.9 per cent in 2008 for urban residents and from 67.7 per cent in 1978 to 43.7 per cent in 2008 for rural residents (China National Statistics Bureau, 2009).
(2) The One Child Policy mandates a long birth interval to protect parents' resources from being depleted (Powell & Steelman, 1995; Yang, 2007). As a result, the second-birth interval during 1980–2000 ranged from 3.5 to 5 years and having an additional child is mostly planned (Chen, Robert, Kim, Xinru, & Cui, 2011) In addition, the One Child Policy has capped the fertility rate to a below-replacement level which could further diminish the resource dilution effect.
(3) Childrearing norms, specifically culturally acceptable altruistic practices among parents have been reshaped during the longstanding campaigns of 'quality childrearing' (*you sheng you yu*). Children with a few siblings might still be able to enjoy the same level of nutrition intake at the cost of the family's consumption of other goods.

Whether having siblings affects resource allocation in nutrition intake may also vary by gender of the child. Girls suffer from discriminatory treatment in both prenatal and postnatal periods (Li et al., 2007; Li & Cooney, 1993). The reluctance to invest resources in girls was especially prevalent among older generations (Das Gupta & Bhat, 1997). Evidence shows that boys are more likely to receive breast-feeding, quality food and medical treatment than girls (Li, 2003). At the same time, in addressing the gender inequality in nutrition intake as fertility is falling remarkably, the 'Parity Effect' hypothesis (Das Gupta & Bhat, 1997) argues that fewer children means girls are likely to receive equal care. If girls are treated as equal to boys, the dilution effect of having siblings should also be equal across gender. 'Intensification Effect' hypothesis (Das Gupta & Bhat, 1997), on the other hand, argues that boys are even more treasured because the decline of fertility is faster than the decline of son preference. Concerning the effect of having siblings on nutrition status, 'Intensification Effect' would suggest that boys would not suffer as much from dilution effect as girls. Some findings on centre-based childcare enrolment suggest that if family resources are scarce, parents often invest more in the eldest son regardless of the gender of his sibling(s) (Zhai & Gao, 2010).

Furthermore, gendered ideal body shape, which encourages girls to be thin, could potentially legitimise less resource allocation in nutrition for a girl, particularly if she has a younger sibling. Women in China are under much greater pressure to lose weight than men (Luo, Parish, William & Laumann, 2005) as the ideal of a thin body type—implying beauty, health and self-discipline—has spread from Western countries to Asia (Cash & Pruzinsky, 2002; Watts, 2002; Wong, Bennink, Wang, & Yamamoto, 2000).

Previous studies on the impact of the number of siblings on child nutrition status in China have yielded mixed findings. Number of siblings is positively associated with the risk of being underweight for children ages 2–6 in rural areas (Brauw & Mu, 2011). No difference in underweight status between children with siblings and children without

siblings was found in Zhejiang, China in 1999 (Hesketh et al., 2003). Regarding obesity, some studies found that being an only child is associated with a higher risk of being overweight in China and some other Asian countries (Chamratrithirong, Singhadej, & Yoddumern-Attig, 1987; Hesketh et al., 2003; Parsons, Power, Logan, & Summerbell, 1999), while some studies did not observe such an association (Yang, 2007) These aforementioned studies mostly identified associations instead of causalities; and the methodological variation might have contributed to the observed mixed findings. We are not aware of any study that attempts to identify the causal impact of having siblings on undernutrition and obesity.

Setting

This study is conducted in the context of China's One Child Policy regime. This unique setting in China provides an opportunity to identify the impact of having younger siblings on child nutrition status in the low fertility era. We exploit a policy variable, monetary fine level, for unsanctioned births as instrument variable to achieve this goal. Background information on the One Child Policy in the following paragraphs helps to explain the method employed in this study.

The One Child Policy has undergone decentralisation and relaxation since 1984 (Greenhalgh, 1986). The localisation of the national policy was a response to China's highly heterogeneous demographic and social conditions, and was designed to facilitate practical policy implementation (Gu et al., 2007). The regional variation of policy-sanctioned number of children per couple varies by regional economic conditions, population density, resistance, as well as minority composition, etc. For example, resistance in poor rural areas is especially high, therefore a second child is allowed under certain conditions (Greenhalgh, 1986). Gu et al. (2007) calculated the policy fertility levels across regions and categorised three groups as of the late 1990s: (1) 'one-child policy' in Beijing, Tianjin, Shanghai, Chongqing, Jiangsu and Sichuan where fertility ranges from 1.06 to 1.27; (2) '1.5-children' policy in 19 provinces where rural residents may obtain a permit to have a second child if the firstborn is a girl. The fertility level varies in these areas from 1.38 to 1.67; (3) 'two-children' policy in five provinces, Hainan, Ningxia, Qinghai, Yunnan and Xinjiang where minorities make up the majority of the population and the fertility rate is 2.01 to 2.37.

The fact that the number of policy-sanctioned children per family is not randomly assigned but related to regional characteristics makes it less than ideal as an instrument variable. Regional characteristics, themselves, can be directly related to child nutrition status. For example, fast food restaurants are more densely located in more developed regions, and rural residents are more likely to be less informed about optimal nutrition status and healthy feeding practice.

The One Child Policy is a complex system that provides for compulsory abortion, reduction of land allotment, demotions if working in the public system, denial of public services for the child and monetary fines for violators. Fine levels vary by location and time, for example, Heilongjiang levied a one-time monetary fine of 120 per cent of annual income in 1983, but in 1989 the fine was raised to 10 per cent of income per year for 14 years (Scharping, 2003).

How have the birth quota and strength of enforcement changed over time? Since the 1990s, compulsory abortion and sterilisation have been gradually abandoned as growing

concern about the social, political, physical and economic consequences of these crude enforcement methods spread (Merli & Smith, 2002). However, there is no reason to believe that enforcement was relaxed. In 1991, adoption of the 'cadre responsibility for family planning system' (*yi piao fou jue*) further strengthened enforcement. Under the cadre responsibility system, the cadres' level of remuneration and their tenure in office and opportunity for promotion are determined by how well their communities comply with birth limits set by officials higher up in the family planning system. In 2000, the 'three unchangeable (*san bu bian*),' an official parlance reinforces: (1) no change of the present policy, (2) nor the birth limits, (3) nor the cadre responsibility system (Merli & Smith, 2002).

However, China's transformation from a centrally planned economy to one dominated by the marketplace had an impact on the family planning system. Since the 1990s, the central government began to retreat from funding local family planning offices. One major strategy adopted by the local offices was to increase fines for non-compliance (Merli & Smith, 2002). Therefore, whether the change of provincial monetary fine level is exogenous to the fertility level or other characteristics that could be related to child nutrition status might become a concern. We will address this concern in the next section.

Data and methods

We draw data from CHNS waves 1991, 1993, 1997, 2000, 2002, 2004 and 2006. These surveys use a multistage, random cluster process to create samples in nine provinces that vary substantially in geography, economic development, public resources, and health indicators (for more details, see Popkin, Du, Zhai, & Zhang, 2010). Like many longitudinal data, CHNS data is also subject to attrition. A close check shows that body mass index (BMI) in the previous wave is not related to the attrition status conditional on a set of observables, suggesting that the attrition is random (See, Table A1). There are 4,293 observations of the eldest children with non-missing values for the main model estimation. We dropped nine observations with BMI values greater than 50 or less than 10 and obtained an effective sample size of 4,284. Then we checked to see if missing values were related to mother's BMI; results showed that this is random (See, Table B2).

Measurement

Dependent variables

One dependent variable is child overweight/obesity, a binary variable (1 = overweight/obesity, 0 = others) constructed by weight and height measured by trained survey team members. Child overweight/obesity is defined using a composite scale based on the Working Group of Obesity in China (WGOC) reference and the International Obesity Task Force (IOTF) reference. For children ages 7–18, we used the WGOC Body Mass Index (BMI) cut-off. For children ages 2–6, for whom a WGOC reference is lacking, we use the IOTF BMI reference. The IOTF reference defines overweight based on data from six countries/regions, including Singapore and Hong Kong. The WGOC reference is based on Chinese children and adolescents. Prior studies indicate that the WGOC reference is the most relevant to the metabolic syndrome for Chinese children and adolescents.

Another dependent variable is child underweight, a binary variable (1 = underweight, 0 = others) measured by IOTF reference.

We did not choose BMI as the dependent variable due to the following considerations. In this paper, we focus on studying the impact of having any younger sibling (s) on first-born children's overweight and underweight status. At different development stages for children under 18, BMI is differently associated with clinical risk factors of cardiovascular disease such as hyperlipidemia, elevated insulin and high blood pressure (Dietz, 2005); therefore BMI cut-offs for overweight and obesity must be age and gender specific.

Measurements of underweight include weight for age, weight for height, height for age, and BMI for age. Among these measurements, BMI for age has been recognised as the most encompassing measurement because it makes use of the information of height, weight and age (Cole, Flegal, Nicholls, & Jackson, 2007). The advantage of BMI for age, compared to weight for height is that it recognises that the weight-height relationship varies by age. In infancy, the ratio of weight/height is larger compared to mid-childhood because this is the period when weight grows fastest relative to height; whereas in later adolescence, as weight continues to grow but height growth stops, the ratio increases again (Cole, 1986).

Main explanatory variable and instrument variable

The main explanatory variable used is whether or not there are any younger sibling(s). This is instrumented using monetary fine level for an unsanctioned birth which varies by year and location. The rationale behind the instrument variable is that higher fine level could reduce the occurrence of unsanctioned birth through financial strain, thus can be used to predict the probability that a first born has any younger siblings. Other potential instrument variables considered include 'whether the couple is allowed to have a second child if the first is a girl'. However, the community (neighbourhood) level policy-permitted number of children per couple (birth quota) is also linked to child nutrition status through the local economic, political and cultural environment. Thus exclusion restrictions criteria do not hold for birth quota. Based on these considerations, birth quota is controlled in the models, instead of being used as an instrument variable.

To measure the total amount of monetary fines parents believe they will incur if they have an unsanctioned birth, we consider four measures based on the information of the mean length of second-birth intervals and the provincial fines levied on unsanctioned births each year. The mean length of second-birth intervals ranged from 3.5 to 5 years from 1975 through 2005 (Chen et al., 2011). So the first measure of perceived fine level we have considered is the fine at the fifth year since a first child is born; and the second measure considered is the fine level at the third year since a first child is born. The third measure is the ten-year average fine since the birth of a first child. The fourth measure is the seven-year average fine level since the first child was born. Because the first two measures only use one year of information, it does not apply to the children born within one or two years since the first child was born. Using multiple-year average fine level allows the instrument variable to be a more reasonable predictor for unsanctioned birth during the multiple years since the first child was born. Compare to a single year fine level, the average fine level also better reflects the expected/perceived fine by the parents. Therefore these multi-year measures were chosen.

We obtained records of provincial fines from 1979 to 2000 (See, Table C3) collected by Scharping (Ebenstein, 2011; Scharping, 2003). Monetary fine is levied as a percentage of annual household income. To calculate the perceived fine levels, we first calculated the present value of the fine for each year in each province. For example, if the fine in 1980 is ten per cent of household income for 14 years, a present value of 1.2283 years of income is calculated for an unsanctioned birth in 1980, with a two per cent discount rate. Then we average the present value of the fine for each year in each province through seven and ten years, respectively, to obtain two measures of perceived fine level.

Other explanatory variables

Other explanatory variables include individual level characteristics: gender (1 = boy, 0 = girl), age, minority status (1-minority, 0 = Han), survey wave, residency (1 = urban, 0 = rural), per capita household income (adjusted by 2006 Consumer Price Index), parental education (1 = high school or above, 0 = lower than high school), and parental height (cm), and community level characteristics: community-level-permitted number of children per couple, average per capita family income, average parental height, percentage of parents holding a high school diploma and community children's gender ratio.

Methods

Maximum likelihood bivariate probit (BP) models (Greene, 1998; Heckman, 1978) correcting for clustering at the individual level are used to identify the impact of having siblings on a child's risk of being overweight and underweight in the low fertility era. Linear instrument variable models are not chosen when overweight and underweight are the outcome variables because in the case that the outcome variable and the endogenous predictor are both binary variables, maximum likelihood bivariate probit models tend to perform better than linear IV models; this is especially true for smaller sample sizes (below 5,000) when the model specification includes additional covariates (Chiburis, Das, & Lokshin, 2011). In addition, when the instrument is weak, two-stage IV model could be seriously biased (Bound, Jaeger, & Baker, 1995).

Models control child's demographic variables age, gender, minority status and family socioeconomic status. Community fixed effect is controlled to capture the time-invariant community characteristics that could be related to the general fine level and simultaneously affect the outcome of interest, such as general socioeconomic development, political environment, traditional value and son preference fixed within the community. Year fixed effect is controlled to capture the national trends over years that might be related to the change of fine levels and child obesity as well. Community-level-permitted number of children per couple, average per capita family income, average parental height, percentage of parents holding a high school diploma and community children's gender ratio are controlled to capture the time-varying characteristics that might be related to the change in fine levels and child nutrition status.

The equations for BP models are set up, where Y denotes outcome variable overweight/obesity, or underweight; S denotes whether having younger siblings, the explanatory variable of interest; Z denotes the average fine level since the first child was born as the instrument variable, and X is a vector of covariates including child's age, gender, minority status, family income adjusted by CPI, urban/rural residency, parental education, parental age,

parental height, community-level average family income, community-level percentage of boy among children, community-level percentage of parents holding a high school diploma, community-level parents' height, community-level allowed number of children per family, community fixed effect and year fixed effect.

$$S_i = 1[\alpha_{10} + \beta_{11}Z_i + \beta_{12}X_i > \xi_{1i}] \tag{1}$$

$$Y_i = 1[\alpha_{20} + \beta_{21}S_i + \beta_{22}X_i > \xi_{2i}] \tag{2}$$

Error terms ξ_{1i} and ξ_{2We} jointly distributed as standard bivariate normal with correlation ρ. The joint probability of ($P_i = 1$, $Y_i = 1$) follows bivariate cumulative distribution and bivariate probit models estimate the parameters by maximising the joint log-likelihood of the two jointly determined variables. ξ_{1i} and ξ_{2i} contain common components such as preference/taste, informal social connections or unobserved wealth and health endowment that affect both having younger siblings and child nutrition status. If $\rho = 0$, then S_i is exogenous after taking into account the influence of the set of covariates. In such cases, the results from univariate probit models and bivariate probit models should be qualitatively the same, and the model can be simplified to a univariate probit model. If ρ is different than 0, a univariate probit model is subject to omitted variable bias. To test this exogeneity hypothesis, likelihood ratio test (Wald test) (Greene, 1998; 2000) was conducted. The ratio of the log likelihood for the bivariate probit model versus the sum of the log likelihood of the two univarite probit models, follows chi-square distribution with one degree of freedom under the null hypothesis $\rho = 0$.

In addition, to examine the proximate mechanisms, we also estimate two-stage linear least squares models to identify if having sibling(s) affects nutrition intake measured as total caloric intake, fat intake and protein intake as well as percentage of calories from fat and protein. The model specification is listed below.

$$S_i = \mu_{10} + \pi_{11}Z_i + \pi_{12}X_i + \varepsilon_{1i} \tag{3}$$

$$Y_i = \mu_{20} + \pi_{21}S_i + \pi_{22}X_i + \varepsilon_{2i} \tag{4}$$

$$\mathrm{Cov}(\varepsilon_{1i,}\varepsilon_{2i}) \sim= 0$$

Is fine level a good instrument? Ideally, fine level only affects a child's weight status through the size of the child's younger siblings after controlling for all community-level effect and national trends. However, having unsanctioned births usually means loss of a portion of disposable income which exacerbates the resource dilution effect on child nutrition status. The treatment effect is the sum of loss of income and resource dilution alleviated by various factors, which is the effect of having younger sibling(s) under the One Child Policy regime. Any significant effect found could be attributed to the combined effects of resource dilution and monetary loss. If there is no significant effect, it would be conservative to conclude the resource dilution effect does not exist.

Is the change in level of fines exogenous? As discussed previously, the general increase of fine level was driven by revenue-generating incentives since the central government stopped funding local family planning offices. The revenue-generating incentive might be related to local economic conditions. If change in fine level is related to local economic conditions, then the validity of the instrument variable is compromised. In addition, the validity of the instrument could also be threatened if the change in fine level is responsive

to the community-level fertility rate. To address these concerns, we examined the change in fine levels from 1991 to 2000 to see if it was a response to the local economic conditions or the previous fertility level in 1991. Results show that after adjusting a set of community-level characteristics, neither the community-level average number of children nor the average per capita income in 1991 predicts the change in provincial fine level from 1991 to 2000 (See, Table D4). Fines were also strictly assessed. Family planning officials reported that about 90 per cent of families who violated the birth quota actually paid the penalty in the 1991 and 1993 waves where these questions were asked.

For first-order girls, one more concern is that the validity of fines could be threatened by parents' under-reporting or other measures taken due to the preference for a son. China observes a gender imbalance at birth and it is arguably a result of underreporting or non-registration and prenatal/neonatal discrimination (Ebenstein, 2011; Hesketh et al., 2005; Merli & Smith, 2002). The sex ratio at birth has been increasing since 1980s, from 108.5 boys per 100 girls in 1982, 113.8 in 1989 (Gu & Roy, 1995), to 121.18 in 2004 (SSBC, 2005). Fine level has been found to causally increase the sex ratio (Ebenstein, 20011). CHNS data is collected by China's Centre of Disease Control, so it is possible that respondents hide first-born girls from the government interviewers, and the probability of a first-born girl being observed (or being reported in the survey) depends on a couple's preference for a son or daughter and a high or low fine level. For example, when the fine level is low and son preference is low, the probability of first-born girls being observed is the highest; whereas when the fine level is high and son preference is high, the probability of a girl being observed is the lowest.

Below we consider two scenarios. In the first scenario, assuming in the population the community level son preference is not related to the level of fine, that is, the communities facing high fine regime and the communities under low fine regime have the same level of average son preference. Then in the community with high fine level, the parents who have above-average-level son preference might be more likely to underreport their first-born girls than the community facing low fine level as a response to the higher fine, therefore, in the high-fine communities, for the girls observed in the sample, the average level of their parents' son preference should be lower than the observed girls in low-fine community. In such case, son preference might be negatively related to fine level among the observed girls. The second scenario assumes that son preference is not randomly distributed among communities, for example, the high fine communities might have higher son preference than the low fine community. In such case, how community level son preference and fine level are related in the sample would be uncertain.

Both scenarios suggest that fine levels could be related to son preference. And son preference together with poverty affects girls' nutrition and health, resulting in marked gender disparity in height and morbidity (Burgess & Zhuang, 2000; Graham, Larsen, & Xu, 1998). Therefore the validity of the instrument for the girls' sample could be undermined. In order to mitigate this potential problem, we control the determinants of son preference at the community level and the individual level. We control residency type because urbanisation and industrialisation are negatively related to son preference (Murphy, Tao, & Lu, 2011). Community-level patrilineal norm (Murphy et al., 2011) is controlled by community fixed effect and time-varying community child gender ratio. Individual-level determinants of son preference such as parental education level and age (Chuang, 1985; Li & Lavely, 2003; Murphy et al., 2011; Yan, 2003) are also controlled. In

analysis, we first control the set of community-level determinants of son preference and then control individual-level determinants of son preference to see if adding these controls makes a difference in the estimates.

Results

Descriptive analyses

Descriptive analysis on main variables of interest by survey year is presented in Table 1. The proportion of boys in the first-born children and adolescents samples has increased over the years, consistent with previous studies on all-order children (Gu & Roy, 1995; SSBC, 2005). The average age in this sample is 11 to 12 before 2000, but increased to 15 and 16 in 2004 and 2006. This is because observations have to be born in 1991 or before to have available values on 10-year average fine levels after they were born. The prevalence of overweight/obesity among this sample increased from around 7.0 per cent in 1991 to 13.3 per cent in 2006. The prevalence of underweight remained about the same, from 5.1 per cent to 6.5 per cent. The proportion having siblings steadily declined from 50.6 per cent to 27.3 per cent. Annual family income steadily increased from 10,100 Yuan to 26,300 Yuan. Urban firstborn children make up 31.0 per cent of the sample in 1991 and 46.8 per cent in 2006, a larger portion compared to all-order children because most of the second-born children are rural residents. Percentage of parents holding high school diplomas has increased over time as has average parental height. The prevalence of children subject to the 1.5-child policy declined over years, so did the prevalence of children subject to the two-child policy. The mean of the 10-year average fine level after the respondent was born increased steadily from 1.23 years of annual family income in 1991 to 2.15 in 2006. Percentage of ethnic minorities among the first-born sample declined over the years. Total daily energy intake remained at a similar level over years, but daily protein intake and fat intake increased.

Having younger siblings and weight status

As shown in panel A of Table 2, we estimated the OLS model (Model 1) and the BP model (Model 2) using the sample of first-born children aged 2–18 using overweight/obesity as the dependent variable, correcting clustering at the individual level. We also explored whether the estimates differ by gender (See Model 3 and Model 4). These models do not include individual level son preference determinants parental age and parents' education. In the OLS model, the estimated coefficient on the younger sibling variable is insignificant. Age is negatively related to overweight/obesity. Family income is positively associated with overweight/obesity.

The results from the maximum-likelihood bivariate models (Models 2–4) suggest that the 10-year average provincial fine level strongly predicts the chance of having younger siblings for the firstborn children's sample ($t = 5.73$), firstborn boys' sample ($t = 4.22$) and firstborn girls' sample ($t = 4.28$). We also estimated the models using seven-year average provincial fine levels as instrument variable, but the results show that seven-year average fine levels are only weakly related to having siblings after controlling for covariates, so it is not used as valid instrument variable here. The results from the BP models also show that the correlation between the error terms of the two equations significantly

Table 1. Descriptive statistics for first-born children ages 2–18 with no missing values in major variables, China Health and Nutrition Survey 1991–2006.

	1991		1993		1997		2000		2004		2006	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Male	.476	.500	.499	.500	.514	.500	.519	.500	.527	.499	.514	.499
Age (years)	11.2	4.94	11.0	4.49	12.36	3.50	13.7	2.55	15.78	1.67	16.8	1.12
Overweight/Obese	.070	.256	.087	.282	.078	.268	.091	.288	.116	.308	.133	.340
Underweight	.051	.219	.058	.234	.065	.247	.058	.234	.052	.221	.062	.240
Have younger sibling(s)	.506	.500	.511	.500	.486	.500	.366	.482	.333	.472	.273	.446
Family real income (in thousand Yuan)	10.1	7.11	11.4	10.0	13.7	10.5	16.1	12.5	21.6	19.8	26.8	28.5
Urban resident	.310	.462	.356	.479	.339	.473	.396	.489	.412	.493	.468	.500
Father high school	.233	.423	.278	.448	.306	.462	.347	.477	.303	.460	.403	.492
Mother high school	.156	.363	.207	.406	.225	.418	.269	.443	.267	.431	.248	.433
Father's height (cm)	166	6.36	166	6.13	167	6.08	168	6.28	168	7.12	168	10.7
Mother's height (cm)	155	5.68	156	5.56	156	5.48	157	5.80	157	8.31	157	9.82
Allow 1.5 children	.420	.494	.334	.471	.350	.477	.350	.477	.430	.495	.316	.466
Allow two children	.189	.390	.112	.316	.220	.414	.229	.420	.038	.192	.015	.121
Seven year average fine	1.01	.621	1.24	.650	1.39	.501	1.67	.688	1.98	.762	2.01	.834
Ten year average fine	1.23	.719	1.45	.749	1.53	.510	1.93	.769	2.17	.835	2.15	.906
Minority	.162	.368	.149	.355	.127	.333	.137	.343	.103	.304	.107	.310
Daily energy intake (1000 kcal)	2.13	.773	2.02	.763	1.98	.630	2.10	.794	2.20	.746	2.03	.626
Number of obs.	1122		1159		881		634		330		158	

Table 2. Results for overweight/obesity from OLS and bivariate probit models for first-born children ages 2–18, CHNS 1991–2006, clustering correction at the individual level.

Overweight/obesity	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A				
	Model 1	Model 2	Model 3	Model 4
Having younger siblings	**−.003(.009)**	**−.177(.322)**	**.103(.641)**	**−.363(.277)**
Age	−.008(.001)***	−.055(.011)***	−.064(.016)***	.003(.005)
Boy	.009(.008)	.013(.071)	N/A	N/A
Family income logged	.011(.005)**	.014(.013)	.024(.006)***	−.002(.012)
Community PB	−.101(.069)	−.200(.134)	−.208(.104)*	.024(.123)
Allow 2 children	.013(.017)	.002(.020)	.024(.026)	−.006(.039)
Allow 1.5 children	−.022(.018)	−.007(.023)	−.041(.027)	−.012(.021)
Marginal effect of IV		**−.083(.018)*****	**−.094(.023)*****	**−.076(.018)*****
Correlation of errors		.100(.052)**	−.123(.077)**	−.140(.076)**
P value: rho = 0		.019	.030	.047
Marginal effect of younger siblings		**−.010(.009)**	**.004(.016)**	**−.021(.030)**
Panel B: adjusting parental age and education				
	Model 5	Model 6	Model 7	Model 8
Having younger siblings	**−.003(.010)**	**−.184(.414)**	**.101(.687)**	**−.356(.278)**
Age	−.007(.001)***	−.049(.013)***	−.065(.018)***	.003(.006)
Boy	.009(.009)	.013(.083)	N/A	N/A
Family income logged	.011(.006)*	.013(.015)	.024(.008)***	−.002(.011)
Community PB	−.095(.067)	−.133(.091)	−.178(.104)*	.028(.122)
Allow 2 children	.010(.023)	.002(.023)	.024(.027)	−.004(.037)
Allow 1.5 children	−.017(.017)	−.007(.022)	−.042(.026)	−.013(.024)
Marginal effect of IV		**−.082(.019)*****	**−.093(.025)*****	**−.076(.019)*****
Correlation of errors		.101(.054)**	−.134(.077)**	−.142(.073)**
P value: rho = 0		.021	.034	.043
Marginal effect of younger siblings		**−.010(.009)**	**.004(.013)**	**−.020(.031)**
Number of observations	4284	4284	2155	2129

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.

different than zero (rho∼ = 0), which suggests that there are unobserved characteristics related to both nutrition status and siblings that OLS or ordinary probit models would fail to control.

The estimates from the BP models (Models 2–4) show that having younger siblings does not predict the risk of being overweight/obese. Models 3 and 4 show that family income increases boys' risk of being overweight/obese but does not affect girls' chance of being overweight/obese. After adjusting for individual-level son preference determinants including parental age and education level, there is little change in the results (See Panel B of Table 2), suggesting the bias that could come from uncontrolled son preference might be small, if it exists at all.

Then using underweight as the dependent variable, we estimated OLS model and BP models on the sample of first-born children aged 2–18 (See Panel A of Table 3). Again, these models do not include individual level son preference determinants such as parental age and education. OLS estimates show that having younger siblings does not affect the first-born child's risk of underweight. The Wald test on the endogeneity of

Table 3. Results for underweight from OLS and bivariate probit models for first-born children aged 2–18, CHNS 1991–2006, cluster at individual level.

Underweight	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A				
	Model 9	Model 10	Model 11	Model 12
Having any younger siblings	**.015(.013)**	**.112(.245)**	**.087(.235)**	**.248(.179)**
Age	.001(.001)	−.006(.004)	.006(.004)	−.016(.006)
Boy	.024(.019)	.073(.047)		
Family income logged	−.011(.007)	−.008(.007)	.002(.009)	−.024(.011)**
Community PB	.069(.068)	.035(.076)	.011(.106)	.011(.134)
Allow 2 children	−.023(.030)	.037(.028)	.008(.039)	.051(.041)
Allow 1.5 children	.039(.025)	.013(.021)	−.004(.031)	.054(.030)
Marginal effect of IV		**−.084(.019)*** **	**−.095(.023)*** **	**−.077(.018)*** **
Correlation of errors	N/A	.300(.045)***	0.017(.022)	−.436(.122)***
P value: rho = 0		.000	.129	.000
Marginal effect of younger siblings		**.011(.015)**	**.007(.013)**	**.026(.019)**
Panel B: adjusting parental age and education				
	Model 13	Model 14	Model 15	Model 16
Having any younger siblings	**.013(.014)**	**.113(.144)**	**.081(.246)**	**.230(.181)**
Age	.001(.001)	−.006(.004)	.006(.004)	−.016(.006)
Boy	.022(.016)	.071(.046)		
Family income logged	−.009(.007)	−.008(.007)	.002(.009)	−.024(.011)**
Community PB	.066(.069)	.035(.076)	.011(.106)	.011(.134)
Allow 2 children	−.020(.032)	.035(.029)	.008(.039)	.051(.041)
Allow 1.5 children	.037(.023)	.011(.022)	−.004(.031)	.054(.030)
Marginal effect of IV		**−.082(.019)*** **	**−.093(.025)*** **	**−.076(.019)*** **
Correlation of errors	N/A	.299(.043)***	0.017(.022)	−.436(.122)***
P value: rho = 0		.000	.136	.000
Marginal effect of younger siblings		**.010(.017)**	**.007(.015)**	**.024(.018)**
Number of observations	4284	4284	2155	2129

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.

having younger siblings suggests that the OLS model is biased by omitted variables. The estimate using BP models is different from OLS model but also shows that having younger sibling(s) does not have significant effect on underweight. Model 12 suggests that family income reduces the risk of underweight only for girls. After adjusting individual-level son preference determinants including parental age and education level, there is little change in the estimates (See, Panel B of Table 3). For Shandong and Guangxi, the level of fines as a proportion of income did not change, therefore its variation over time was driven by income growth. We excluded these two provinces and re-estimated the models but did not observe significant material change on the effects (Table E5).

To further explore the role potentially played by son preference, we divided the sample by one of the most important indicators of son preference: the type of residence (Murphy et al., 2011; Yan, 2003). Results are shown in Table 4 and indicate that in urban areas with low son preference (Chuang, 1985; Li & Lavely, 2003), there is no effect of having siblings on a child's underweight status. In rural areas where son preference is supposedly higher, the effect on a girls' underweight status still falls short of significance.

Table 4. Results for underweight from OLS and bivariate probit models for first-born children ages 2–18 by residence type, CHNS 1991–2006, cluster at the individual level.

Underweight	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A: Urban children				
	Model 17	Model 18	Model 19	Model 20
Having any younger siblings	**−.021(.016)**	**−.101(.103)**	**−.170(.224)**	**−.059(.199)**
Age	.004(.009)	−.011(.006)	.005(.005)	−.019(.011)
Boy	.022(.017)	.077(.054)		
Family income logged	−.009(.008)	−.012(.009)	.002(.009)	**−.022(.011)****
Community PB	.070(.066)	.033(.065)	.045(.129)	.025(.104)
Allow 2 children	−.028(.024)	−.034(.029)	.001(.007)	−.040(.051)
Allow 1.5 children	.011(.014)	.033(.027)	−.002(.020)	.051(.041)
Marginal effect of IV		**−.091(.039)*****	**−.099(.040)*****	**−.087(.029)*****
Correlation of errors	N/A	−.033(.037)	−.011(.024)	−.040(.101)
P value: rho = 0		.221	.389	.206
Marginal effect of having siblings		**−.013(.014)**	**−.015(.022)**	**−.005(.017)**
Sample size	1469	1469	740	729
Panel B: Rural children				
	Model 21	Model 22	Model 23	Model 24
Having any younger siblings	**.015(.009)***	**.203(.146)**	**−.009(.009)**	**.322(189)**
Age	.002(.002)	.007(.004)	.008(.005)	−.021(.022)
Boy	.009(.008)	.089(.056)		
Family income logged	−.003(.005)	−.006(.006)	.004(.010)	**−.028(.013)****
Community PB	.014(.033)	.036(.086)	.015(.110)	.017(.141)
Allow 2 children	.012(.014)	−.034(.036)	.011(.069)	.055(.081)
Allow 1.5 children	.010(.010)	.024(.042)	−.007(.044)	.059(.070)
Marginal effect of IV		**−.075(.029)*****	**−.101(.041)*****	**−.046(.020)*****
Correlation of errors	N/A	.140(.051)**	0.007(.011)	.312(.172)*
P value: rho = 0		.001	.209	.015
Marginal effect of younger siblings		**.011(.012)**	**−.001(.013)**	**.027(.019)**
Number of observations	2815	2815	1415	1400

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Parents' height, minority status, parental age and parental high school diploma, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.

To test whether the stage of development makes a difference in resource dilution, we divide the data into two periods - Period 1 which includes 1991, 1993 and 1997 and Period 2 which includes 2000, 2004 and 2006 to estimate OLS model and BP models for these two periods separately (See, Table 5). 1997 was used as the cut-off based on the considerations listed below:

(1) The years from 1997 to 2000 were a landmark period in China's market reforms when intensified country-wide industrial restructuring and privatisation were carried out which set off unprecedented economic growth (Xing & Li, 2012).
(2) In 2001, China was admitted to the World Trade Organisation(WTO) which brought China's market reforms and economic development into a new era (Meng, 2004; Xing & Li, 2012).

Model 25 through Model 32 are estimated using data from Period 1, controlling individual level son preference determinants parental age and education. Again, OLS estimates show that having younger siblings does not affect the first-born child's risk of

overweight or underweight. The Wald test on the endogeneity of having younger siblings suggests that the OLS model is biased by omitted variables. For both boys and girls, the estimates using BP models suggest that having younger sibling(s) does not have significant effect on overweight status. Different from what is observed based on the data from the entire time frame, the results of Model 32 suggest a marginally significant positive effect of having younger sibling(s) on first-born girls' chance of being underweight ($P < 0.1$) (See, Panel B of Table 5). Using the data from Period 2, the instrument variable falls short of significance in predicting the probability of having any younger sibling(s) for both first-born boys and first-born girls due to small sample size (for boys, the number of observations is 580; for girls, 542) (Table F6). No effect was found using OLS models. The contrast of findings suggest that the lack of resource dilution effect for girls through the years from 1991 to 2006 were driven by post-1997 period when favourable effect of economic growth, among others, outperformed resource dilution effect.

Table 5. Results for overweight/obesity from OLS and bivariate probit models for first-born children ages 2–18, CHNS 1991–1997, adjusting parental age and education, clustering correction at the individual level.

	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A: Overweight/Obesity				
Panel A:	Model 25	Model 26	Model 27	Model 28
Having younger siblings	**−.007(.012)**	**−.102(.276)**	**.098(.467)**	**−.298(.187)**
Age	−.006(.001)***	−.051(.014)***	−.067(.019)***	.005(.008)
Boy	.012(.010)	.034(.097)	N/A	N/A
Family income logged	.008(.007)	.019(.017)	.027(.009)***	−.005(.013)
Community PB	−.056(.067)	−.098(.089)	−.167(.101)*	.033(.129)
Allow 2 children	.026(.027)	.006(.054)	.014(.019)	−.004(.028)
Allow 1.5 children	−.001(.015)	−.010(.027)	−.022(.023)	−.012(.026)
Marginal effect of IV		**−.088(.016)*** **	**−.099(.029)*** **	**−.072(.018)*** **
Correlation of errors		.305(.184)*	−.156(.079)**	−.164(.076)**
P value: rho = 0		.056	.040	.023
Marginal effect of younger siblings		**−.014(.034)**	**.011(.039)**	**−.025(.034)**
Number of observations	3162	3162	1575	1587
Panel B: Underweight				
	Model 29	Model 30	Model 31	Model 32
Having any younger sibling	**.023(.034)**	**.136(.130)**	**.090(.276)**	**.276(.161)***
Age	−.005(.019)	−.005(.007)	.007(.006)	−.012(.010)
Boy	−.073(.069)	−.089(.078)		
Family income logged	−.019(.016)	−.009(.011)	.005(.013)	−.029(.014)**
Community PB	.564(.984)	.078(.092)	.017(.103)	.017(.137)
Allow 2 children	.056(.087)	.017(.019)	.012(.040)	.019(.049)
Allow 1.5 children	.023(.020)	.019(.023)	−.001(.025)	.056(.038)
Marginal effect of IV		**−.088(.016)*** **	**−.099(.029)*** **	**−.072(.018)*** **
Correlation of errors	N/A	.350(.063)***	0.027(.014)**	−.422(.118)***
P value: rho = 0		.000	.049	.000
Marginal effect of younger siblings		**.014(.013)**	**.009(.019)**	**.025(.013)***
Number of observations	3162	3162	1575	1587

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.

Notes: Parents' age, parents' holding high school diploma, parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.

Having younger siblings and nutrition intake

To understand the relationship between having younger siblings and risk of malnutrition, we examine the impact of having younger siblings on the first-borns' nutrition intake as shown in Table 6. Wald F statistics for weak instrument all exceeds 10, suggesting that the instrument approach is acceptable. The results from two-stage least squares models show that having younger siblings do not have impact on first-borns' nutrition intake regardless of their gender using the data from 1991–2006 (See Panel A of Table 6). These results are consistent with the findings by Li (2003) and Liang, Gibson, and Stillman (2015). However, using data from 1991 to 1997, the results of the two-stage least squared model (Model 38, See Panel B of Table 6) suggest a marginally significant effect that having younger siblings might reduce the first-born girls' calorie intake by 58.8 Kcal during the period between 1991 to 1997.

Discussion and conclusions

No previous study has identified the causal inference on having younger siblings and child overweight and underweight status under the One Child Policy regime. This study exploits the variation of fine level on unsanctioned birth by location and time to instrument whether the first-borns have any younger sibling so as to identify its impact on child nutrition status. Using China Health and Nutrition Survey 1991, 1993, 1997, 2000, 2004 and

Table 6. Results on daily nutrition intake (kcal) by estimating two-stage instrument variable models for first-born children ages 2–18, CHNS 1991–2006, correcting clustering at the individual level.

	2SLS All first-born children	2SLS First-born boys	2SLS First-born girls
Panel A: Using data from 1991–2006			
	Model 33	Model 34	Model 35
Having any younger sibling	**−32.1(47.7)**	**5.08(44.1)**	**−83.9(88.7)**
Age	104(4.00)***	110(8.34)***	95.4(10.5)***
Boy	78.5(21.1)***		
Family income logged	16.23(4.01)***	20.1(7.22)***	11.8(6.01)*
Community PB	23.0(19.7)	31.4(29.0)	
Allow 2 children	−12.5(20.4)	−15.7(31.9)	−3.45(12.1)
Allow 1.5 children	13.8(43.2)	19.1(78.3)	12.7(33.0)
Wald F statistic for weak instrument	**25.4**	**11.4**	**10.4**
Number of observations	4284	2155	2129
Panel B: Using data from 1991–1997			
	Model 36	Model 37	Model 38
Having any younger sibling	**−33.8(49.9)**	**8.01(53.6)**	**−58.8(34.2)***
Age	101(6.00)***	107(8.38)***	89.0(12.5)***
Boy	72.15(25.1)***		
Family income logged	19.1(5.01)***	22.6(7.33)***	14.9(7.01)**
Community PB	12.0(16.8)	36.9(39.0)	
Allow 2 children	−19.1(23.6)	−11.7(11.1)	−8.99(13.9)
Allow 1.5 children	9.13(43.2)	13.7(71.6)	4.90(38.8)
Wald F statistic for weak instrument	**24.7**	**16.3**	**14.9**
Number of observations	3162	1575	1587

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Parents' age, parents' holding high school diploma, parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys.

2006, we found that under the low fertility era, having younger sibling(s) generally does not affect a firstborn child's risk of being overweight, neither does it increase the risk of being underweight regardless of the gender of the firstborn child. The only exception found was during the period of 1991–1997 when the economy was relatively less developed and the proportion of family income spent on food was higher: having younger sibling(s) might have decreased the first-born girls' nutrition intake and increased their risk of being underweight.

Collectively, these results do not support a strong resource dilution effect under the low fertility era in China for the first-born children these are consistent with the findings that having siblings does not generally affect nutrition intake in the low fertility era in China (Li, 2003; Liang *et al*., 2005; Yang, 2007). Lack of resource dilution effect might be largely due to the low fertility rate, mandated long birth interval, capacity building in planning for an additional birth and parents' altruistic child-raising practice as proposed or discovered in previous studies on quantity-quality trade-off effect in Asian countries (Bisai et al., 2014; Liu, 2015; Millimet & Wang, 2006). This study adds to the literature that in less developed countries, family planning and certain child rearing practice could make considerable difference to offset the resource dilution effect.

However, the marginally significant resource dilution effect found among the first-born girls during the period from 1991 to 1997 shed some light on the importance of stage of economic development as recognised in some previous studies (Li et al., 2008; Liu, 2015). Owing to the rapid economic growth during the years covered in the survey, the Engel's Coefficient (proportion of family income spent on food) decreased drastically. Having more than one child might have diluted the family resource allocated to a girl's basic nutrition needs during the years from 1991 to 1997 to the extent that factors like long birth interval, capacity planning and other family members' sacrifice of their own consumption were not able to offset the dilution effect. This result also suggests that girls might not have received equal resource allocation in this period compared to boys despite the low fertility rate, contradictory to the prediction of the 'Parity Effect' hypothesis (Das Gupta & Bhat, 1997). The unequal treatment could also be due to the gendered body shape ideology (Luo, Parish, & Laumann, 2005) that praises girls for being thin, which could have legitimised less resource allocation of nutrition to girls.

The finding that being an only child does not increase the risk of being overweight contradicts some popular notions that the Little Emperor Syndrome 'is shaping the society in unexpected ways that may culminate into a future behavioural time-bomb' (Branson, 1988). This study suggests that despite the possibility of being granted much more bargaining power in food consumption than the children with siblings, those with only child status were subjected to some measure of effective control which protects them from having higher risk of obesity.

One additional and very interesting finding is that family income increases the risk of being overweight for boys but not girls, whereas family income reduces girl's risk of being underweight but not boys. This contrast might result from the possibility that girls are under greater pressure to keep thin (Luo et al., 2005). Therefore, they do not respond to the increase in access to resources when there is risk of being overweight; but, at the same time, increase in income protects the first-born girls from the risk of being underweight. This does not make much difference to first-born boys, however,

suggesting that boys are protected from being underweight regardless of family income. For first-born boys, although their risk of being overweight responds to family income, it is not affected by the presence or absence of younger siblings. It could be that other family members absorbed this cost. For first-born girls, we observed that their obesity status did not respond to family income; neither did it respond to resource dilution from having a younger sibling.

In conclusion, this study did not find strong resource dilution effect under the low fertility era in China for the first-born children. Since the mean ideal family size has generally declined over the period 1980–2000 to below two children a family (Basten & Gu, 2013), further relaxation of the One Child Policy in China to allow all couples to have two children or even more should not likely pose a significant risk to children's basic nutrition needs in the low fertility era.

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Appendix

Table A1. Logistic regression on attrition status by characteristics at the previous wave, for first-born children age 2–18, CHNS1991–2006 (robust standard error adjusted at personal ID level).

Gender	−0.059(.310)
Age	0.087(.007)***
BMI at previous wave	1.01(.091)
Log family income	.033(.032)
Urban residence	−0.012(.005)**
Father high school or higher	0.071(.027)**
Mother high school or higher	0.131(.005)**
After 1997	0.153(.020)**
Father's height	0.001(.203)
Mother's height	0.012(.004)**
Pseudo R2	0.139
N	4284

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Province fixed effects are controlled.

Table B2. Regress mother's BMI on Missing status for first born children aged 2–18, CHNS 1991 to 2006, correcting clustering at individual level.

Mother's BMI	Coefficient
Missing	1.34(1.15)
Age	.018(.117)
Gender	1.58(.98)
Urban residency	.663(.105)***
R squared	.002
N of observations	7910

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.
Notes: Province fixed effects are controlled.

Table C3. Monetary punishments for excess fertility, China 1979–2000.

Province	First report	Second report	Third report	Fourth report	Fifth report
Liaoning	1979: 14Y, 10%	1980: 14Y, 10%	1988: 14Y, 10%	1992: 1Y, 500%	1997: 1Y, 500%
Heilongjiang	1982: 14Y, 10%	1983: 1Y, 120%	1989: 14Y, 10%		
Jiangsu	1982: 10Y, 10%	1990: 1Y, 300%	1995: 1Y, 300%	1997:1Y, 300%	
Shandong	1996: 1Y, 100%				
Henan	1982: 7Y, 15%	1985: 7Y, 15%	1990: 7Y, 30%		
Hubei	1979: 14Y, 10%	1987: 5Y, 10%	1991: 5Y, 60%	1997: 5Y, 60%	
Hunan	1979: 14Y, 5%	1982: 5Y, 10%	1989: 1Y, 200%		
Guangxi	1994: 1Y, 500%				
Guizhou	1984: 14Y, 10%	1998: 1Y, 500%			

Notes: Taken from Ebenstein (2011). Monetary punishment listed above as 'Year of report: length of wage deduction, per cent of annual salary'. Fines that are levied as one-time punishments are listed above as being collected in a single year.

Table D4. Regress change of fine level from 1991 to 2000 on 1991 community level characteristics, correcting clustering at individual level.

Community level characteristics at 1991	Change of fine level from 1991 to 2000
Average number of children per family	−.237(.209)
Average per capita family real income	4.68e-06(.0000106)
Percentage of boys among children	−5.38 (.467)***
Percentage of minority	2.70 (.157)***
Two-child zone	.591(.130)***
1.5-child zone	1.28(.094)***
Percentage of fathers holding high school diploma	−.457(.384)
Percentage of mothers holding high school diploma	1.31 (.441)***
Average father's height	.136(.019)***
Average mother's height	−.010(.004)
R-squared	0.4201
Number of observations	2152

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.

Table E5. Results for overweight/obesity from OLS and bivariate probit models for first-born children ages 2–18, CHNS 1991–2006, excluding Shandong and Guangxi, adjusting parental age and education, clustering correction at the individual level.

	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A: Overweight/obesity				
	Model 39	Model 40	Model 41	Model 42
Having younger siblings	**−.003(.012)**	**−.161(.500)**	**.131(.600)**	**−.301(.209)**
Age	−.007(.001)***	−.043(.011)***	−.063(.014)***	.004(.006)
Boy	.009 (.008)	.015(.079)	N/A	N/A
Family income logged	.012(.007)*	.014(.013)	.022(.007)***	−.004(.012)
Community PB	−.091(.070)	−.123(.093)	−.180(.109)*	.026(.128)
Allow 2 children	.009(.021)	.002(.027)	.027(.029)	−.005(.039)
Allow 1.5 children	−.011(.014)	−.006 (.012)	−.039(.024)	−.016(.023)
Marginal effect of IV		**−.093(.017)*** **	**−.103(.07)*** **	**−.089(.023)*** **
Correlation of errors		.106 (.057)*	−.129 (.067)*	−.156 (.085)*
P value: rho = 0		.022	.045	.056
Marginal effect of younger siblings		**−.010(.008)**	**.007(.016)**	**−.015(.029)**
Number of observations	3309	3309	1659	1650
Panel B: Underweight				
	Model 43	Model 44	Model 45	Model 46
Having any younger sibling	**.013(.014)**	**.103(.112)**	**.004(.459)**	**.241(.192)**
Age	.001(.001)	−.006(.004)	.007(.008)	−.016(.006)
Boy	.024(.018)	.075(.049)		
Family income logged	−.010(.007)	−.008(.008)	.012(.013)	−.025 (.012)**
Community PB	.046(.067)	.065(.077)	.021(.156)	.011(.136)
Allow 2 children	−.013 (.033)	.038(.030)	.021(.053)	.055(.044)
Allow 1.5 children	.024 (.033)	.013(.020)	−.009(.041)	.056(.036)
Marginal effect of IV		**−.093(.017)*** **	**−.103(.027)*** **	**−.089(.023)*** **
Correlation of errors	N/A	.301 (.042)***	.022(.019)	−.455 (.132)***
P value: rho = 0		.000	.139	.000
Marginal effect of younger siblings		**.009(.016)**	**.010(.017)**	**.025(.019)**
Number of observations	3309	3309	1659	1650

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.

Notes: Parents' age, parents' holding high school diploma, parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.

Table F6. Results for overweight/obesity from OLS and bivariate probit models for first-born children ages 2–18, CHNS 2000–2006, adjusting parental age and education, clustering correction at the individual level.

	OLS All first-borns	Bi-Probit All first-borns	Bi-Probit First-born boys	Bi-Probit First-born girls
Panel A: Overweight/Obesity				
Panel A:	Model 47	Model 48	Model 49	Model 50
Having younger siblings	**−.010(.078)**	**−.135(.576)**	**−.048(.341)**	**−.147(.284)**
Age	−.005(.002)**	−.017(.011)*	−.031(.018)*	.008(.013)
Boy	.023 (.056)	.056(.107)	N/A	N/A
Family income logged	.008(.007)	.013(.034)	.021(.010)**	−.009(.298)
Community PB	−.034(.090)	−.068(.100)	−.112(.098)	.029(.147)
Allow 2 children	.033(.023)	.023(.067)	.032(.029)	−.001(.014)
Allow 1.5 children	−.014(.057)	−.028(.056)	−.033(.052)	−.022(.043)
Marginal effect of IV		**−.094(.066)**	**−.133(.092)**	**−.066(.048)**
Correlation of errors		.782 (.354)**	−.177(.099)**	−.214 (.039)***
P value: rho = 0		.041	.048	.000
Marginal effect of younger siblings		**−.014(.042)**	**−.003(.139)**	**−.014(.212)**
Number of observations	1122	1122	580	542
Panel B: Underweight				
	Model 51	Model 52	Model 53	Model 54
Having any younger sibling	**−.005(.089)**	**.100(.180)**	**.130(.341)**	**.067(.483)**
Age	−.008(.039)	−.002(.011)	.003(.008)	−.005(.022)
Boy	−.091(.088)	−.078(.076)		
Family income logged	−.023(.056)	−.012 (.014)	.003(.012)	−.013(.008)
Community PB	.238(.230)	.040(.067)	.043(.293)	.028(.087)
Allow 2 children	.078 (.093)	.055(.062)	.045(.066)	.066(.092)
Allow 1.5 children	.011 (.040)	.023(.084)	.003(.046)	.043(.087)
Marginal effect of IV		**−.095(.066)**	**−133(.092)**	**−.066(.048)**
Correlation of errors	N/A	.429 (.088)***	0.009(.005)*	−.238 (.121)**
P value: rho = 0		.000	.058	.047
Marginal effect of younger siblings		**.009(.015)**	**.011(.062)**	**.003(.101)**
Number of observations	1122	1122	580	542

*$P < 0.1$.
**$P < 0.05$.
***$P < 0.01$.

Notes: Parents' age, parents' holding high school diploma, parents' height, rural/urban residency, minority status, community level average income, community average parental height, community percentage of parents holding high school diploma, community fixed effects and year fixed effects are controlled in all models. Community PB is community percentage of boys; Correlation of errors is correlation of the errors of two equations.