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**Do Siblings Take Your Food Away?
Using China's One-Child Policy
to Test for Child Quantity-Quality Trade-Offs**

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Abstract

We test for the existence of a trade-off between child quantity and quality in Chinese families. We use changes over time and space in the local stringency of the one-child policy as a source of exogenous variation in family size. Investment in child quality is measured by intake of three nutrients, using seven waves of data from the China Health and Nutrition Survey. For all three nutrients, a quantity-quality trade-off is apparent, which persists for fats if child-specific effects are introduced. The trade-off would be less apparent if exogenous sources of variation in family size were ignored.

Keywords

child quality
nutrients
one-child policy
quantity-quality trade-off

JEL Classification

I12; J13; O15

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1. INTRODUCTION

For decades, the relationship between the number and the quality of children in a family has been of interest to demographers and economists. The quantity-quality trade-off theory suggests a negative relationship between the two, stating that 'an increase in quality is more expensive if there are more children because the increase has to apply to more units' and 'an increase in quantity is more expensive if the children are of higher quality, because higher-quality children cost more' (Becker and Lewis 1973, p.S280). This theory, plus the evidence that children from larger families tend to have worse outcomes, underpins many interventions in developing countries where fertility rates may be considered too high to enable sustainable economic development (Angrist, Lavy and Schlosser 2010). In particular, it is expected that policies discouraging large families should lead to increased human capital, which will result in higher lifetime earnings and better life outcomes at the individual level and will contribute to faster economic development in the aggregate.

Despite the importance of the quantity-quality trade-off theory, empirical evidence for these relationships is mixed. Li, Zhang and Zhu (2008) point out that the trade-off is more likely in developing countries, where parents bear a larger share of child rearing cost than in developed countries where the welfare state may subsidize these costs. Empirical research is also complicated by the fact that both the number of children and the investment in each child are choice variables. Thus, studies treating family size as exogenous may not uncover causal effects. To get around this problem, many studies seek to exploit exogenous variation in fertility. For example, Rosenzweig and Wolpin (1980), Li *et al.* (2008), Angrist *et al.* (2010) and Abdul-Razak *et al.* (2015) use multiple births, while Conley and Glauber (2006), Lee (2008), and Angrist *et al.* (2010) use the gender composition of early births.¹ One contributor to exogenous variation in family size, which this paper uses, is the changing stringency of family planning policy (Qian 2009, Liu 2014, Huang 2015 and Zhang *et al.* 2016). Another complication is from measures of child quality; sometimes a trade-off is apparent for one measure of child quality but is very weak for another, even with the same data and empirical methods (Black *et al.* 2005 and Liu 2014).

This paper tests for a trade-off between the quantity and quality of children in China. Intakes of three nutrients are used as the measure of child quality. Two approaches are used to get around endogeneity problems and to control unobserved heterogeneity. First, we use the one child policy (hereafter, OCP) in China as a source of exogenous variation in family size in two-stage least squares (2SLS) models. Second, we use fixed effects (FE) models to examine how a change in sibling numbers affects nutrient intakes, allowing for unobserved heterogeneity of children. In other words, the longitudinal data let us examine within-child variation over time, to allow for child-specific responses that could bias cross-sectional results if there are unobservable factors that confound comparisons of outcomes prior to, and

¹ Bhalotra and Clarke (2016) note that twins may not be quasi-random, due to positive selection of healthier women into twinning. Studies that ignore this form of selection may be biased in the direction of not finding a quantity-quality tradeoff even when it exists.

after, a change in family size. A further advantage of the FE approach is that it is not affected by the weakening identification power of the OCP after 1997 as restrictions were gradually relaxed. With results from these two methods in conjunction, we provide a comprehensive test for the existence of a quality-quantity trade-off.

In particular, the differing strength of the OCP over time and space aids the analysis by providing a source of identifying fertility variation. For example, the total fertility rate fell dramatically, from over three before the introduction of the policy to just above 1.5 by 2006.² Amongst children in our estimation sample, only a quarter were from one-child families in 1991 but almost one half were by the year 2000. Also, boys have a higher probability of being the only-child than do girls, in all survey waves. This difference is very likely due to OCP exceptions that allowed families to have a second birth if the first child is a girl, and highlights the policy-induced variation in family structure.

The measures of child quality used in this paper are children's intakes of dietary energy, protein, and fat. Prior studies using the OCP as a source of exogenous variation in family size use education and health outcomes as the indicators of child quality. For example, Liu (2014) uses children's height-for age and Huang (2015) uses enrolments in post-compulsory schooling while earlier studies for China, with different identification approaches, focus on similar health and education outcomes (Li *et al.* 2008, Lu and Treiman 2008 and Rosenzweig and Zhang 2009). Although less studied, nutrient intakes are not only measures of parental investment in children, they are also important inputs into human capital that have their own long-term impacts (Behrman *et al.* 1988). We contend that nutrients directly reflect how resources are allocated to each child, and give a good testing ground for the quantity-quality trade-off theory. Compared to educational and health outcomes, nutrient intakes are less affected by unobserved personal factors, such as ability and genetics. Thus, in conjunction with the exogenous variation over time and space due to differing strength of the OCP, the use of these novel indicators of investment in children may provide more robust tests of the quantity-quality trade-off than in the existing literature.

In testing for a quantity-quality trade-off, we also shed light on factors influencing the fertility level, with the first-stage equation of the 2SLS approach predicting sibling numbers. Building on Liu (2014), in this paper we define a dynamic measure of OCP strength that accounts for the change in local OCP regulations and how the changes affect local women's eligibility to have a second child. This indicator, along with urban *Hukou* status and the local development level, is estimated to be significant influences on fertility levels.

The remainder of the paper is structured as follows: Section 2 provides a brief review of the one-child policy, while Section 3 discusses the empirical methods. Section 4 describes the

² Total fertility rate is 'the number of children that would be born to a woman if she were to live to the end of her childbearing years and bear children in accordance with current age-specific fertility rates' (The World Bank). Appendix 1 shows the total fertility rate of China from 1977 to 2009.

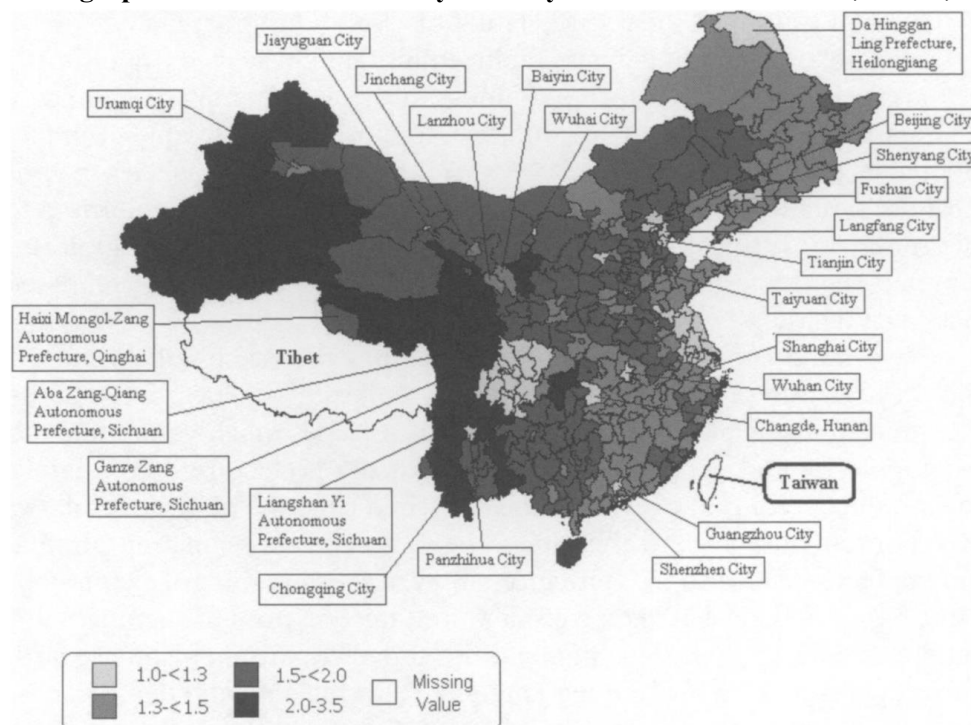
CHNS data we use, paying particular attention to the measures of OCP strength and the nutrient intakes. Section 5 reports the regression results and the discussion and conclusions are in Section 6.

2. BACKGROUND TO THE ONE-CHILD POLICY

The one-child policy was introduced in 1979, and originally limited all couples in China to strictly one child. After four years as a national-level policy, the OCP was decentralized and started to become less restrictive (Greenhalgh 1986), possibly in response to various problems caused by the initial policy, including rapid growth in the number of abortions (Hesketh *et al.* 2005). Starting in 1984, different exceptions to the strict rule of one child per couple were introduced, and these were applied in different areas, and also changed over time.

The decentralized OCP is represented in Figure 1 using the notion of 'policy fertility' as reported by Gu *et al.* (2007). Policy fertility denotes the fertility levels that would obtain locally if married couples had births at exactly the levels permitted by local policy. The data in the figure come from 420 prefecture-level units in China in the late 1990s, with areas given a darker shade representing those where more couples were allowed to have more than one child. There was clearly large variation across the country, with parts of western China having policy fertility rates between 2.0 and 3.5, compared with much of central and southern China where the policy fertility rates were from 1.3 to 2.0. Only small parts of the country (areas with dense populations) had policy fertility rates that were as low as from 1.0 to 1.3.

Figure 1: Geographic Distribution of Policy Fertility at the Prefecture Level, China, late 1990s



Source: Gu, B, Wang, F, Guo, Z and Zhang, E (2007) 'China's Local and National Fertility Policies at the End of the Twentieth Century' *Population and Development Review*, 33(1) pp.143, Figure 2.

Further relaxation and decentralization followed in 1997 with the ‘Opinion of the Ministry of Public Security on Improving the Administrative System of *Hukou* in Rural Areas’. This Ministry controls the *Hukou* registration system and provides identity booklets. The 1997 changes permitted out-of-plan children and the children of unmarried couples to receive birth registration, further reducing effects of family planning policy on fertility (Li et al., 2010). However, some provinces still required the local Public Security Department to be notified of unsanctioned births in order for babies to be registered, or needed similar documentation from the Population and Family Planning Bureau to allow birth registration (Li et al. 2010). Thus it remained true that for many Chinese parents the OCP remained an important exogenous factor affecting their number of children.

In 2002, the Law on Population and Family Planning was adopted and formalised the family planning principle of ‘advocate one-child per couple’ while allowing local exceptions for a second child (The Legislative Affairs Commission of the Standing Committee of the National People’s Congress of the People’s Republic of China, 2002). Under the 2002 law, couples volunteering to have only one child regardless of the local exceptions get a ‘Certificate of Honour for Single-Child Parents’ and are eligible for small cash rewards until their child is a certain age. On the other hand, couples who have more children than local policy allows need to pay a social maintenance fee for the unsanctioned birth(s), which is often a multiple of the previous year’s average household disposable income in the area where they live.³

In this research, we use survey data collected between 1991 and 2009. Over these two decades the strength of the one-child policy varied considerably. Thus we are able to examine impacts on children who had been born in periods when the OCP had been strictly applied, and also when it had been somewhat relaxed, in addition to variation over space in the stringency of the policy. We successfully match each child to a mother, and hence are able to see how the OCP applied to the mother, given her birth history and the status of her and her children. Consequently, the impacts from the OCP on the number of siblings of the sample children are well captured in these survey data.

While it is outside of our sample period, it is worth noting recent changes to China’s family planning policy that let all legally married couples have two children notwithstanding all of the limitations introduced in previous regulations (Xinhua Net 2015). This large break from previous policy heralds the end of the one-child policy but not necessarily the end of

³ The CHNS data show that in the first survey wave (in 1991) almost one quarter of Han children aged less than 18 might be considered as the result of ‘unsanctioned births’ due either to parents being too young or to the number/age of elder siblings of the child suggesting that according to the rules in place when the child was born, they were in excess of the number of children allowed by the local OCP. Some of these apparently unsanctioned births may reflect survey reporting errors, in terms of misdating ages, but also indicate that fines for unsanctioned births may not have been so high as to present an impossible barrier to extra births. The proportion of the sample who appear to be unsanctioned births falls quickly from 1993 onwards.

attempts by policy makers in China to intervene into fertility decisions. In other words, the number of children is still not a completely free choice of the parents, unlike in most of the rest of the world. The official end of the OCP also bookends a generation of children, born between 1980 and 2015, where many of them grew up without any siblings. For example, the share of children from sole-child families in our sample increased from about one-quarter in 1991 to over one-half by 2006, and in urban areas exceeded 70% in that year. Thus it is not exaggerating to think of the OCP as one of the largest policy-induced natural experiments in human history.

3. EMPIRICAL METHODS

We start by regressing nutrient intakes on measures of the number of children, along with other control variables, using Ordinary Least Squares (OLS). We initially ignore the panel structure of the data. Most previous studies do not have panel data so we want to see if the lack of controls for unobservable fixed effects makes prior studies more likely to find a trade-off between quantity and quality that disappears once fixed effects are introduced.

Following the recent literature, our general estimation approach is as seen below:

$$Nut = \alpha_0 + \alpha_1 Sibling + \alpha_2 X + \alpha_3 Z + \varepsilon \quad (1)$$

where Nut is the vector of intakes of energy, protein and fat for each child. The vector X has a set of child characteristics, including gender, age (in quadratic form), whether first-born, *Hukou* status, and province of residence. We use the urban/rural status from the *Hukou* registration (a *de jure* measure) because the OCP restrictions are based on *Hukou* status.⁴ Also, the social benefits system provides differential access to food subsidies, and to education and health facilities, based on registration status rather than on residency. The vector Z has the years of education of the parents, the median nutrient intake levels of adults in each community, the median household per-capita income (in 2009 prices), and an index of local development based on 12 community-level attributes.⁵

The main focus is the coefficient α_1 on the variable *Sibling*, which is the number of siblings the child had at the time of the survey (aged under 18 and living in the same household). This estimates the impact of sibling numbers on the nutrient intake of the child, and negative value would suggest a trade-off between quantity and quality of children.

⁴ About 12 percent of children live in a different sector (urban or rural) than the sector of registration. See Appendix 2 for details.

⁵ The index is provided in the CHNS database and is constructed from community level indicators of: population density, education, sanitation, housing, transportation, communications, health, market development, economic development, diversity, and indicators of modern markets and social services.

However, with OLS this can only be interpreted as a conditional correlation since family size is likely to be endogenous (Black *et al.* 2005, Li *et al.* 2008 and Liu 2014). Therefore, in this research we also use instrumental variables (IV) and fixed effects (FE) approaches to get around this problem.

We use the community level OCP strength as the identifying instrument in the first stage regression for sibling numbers, with the same control variables as in equation (1):

$$Sibling = \beta_0 + \beta_1 OCP + \beta_2 X + \beta_3 Z + \delta \quad (2)$$

The variable *OCP* is the share of women in the community eligible to have two children under the local OCP at any stage up to one year prior to the survey (for women aged from 20 to 49 who were not single). During the period covered by the data, a legally married woman in China could have an extra child if she satisfied the exceptions imposed by the local implementation of the OCP.⁶ Since women can choose to have their second child after meeting the conditions for the exception, we follow Liu (2014) in defining a woman as eligible to have two children if she met the OCP conditions at any stage up to the year prior to the survey. The following four OCP exceptions are included, with consideration given to the birth gap constraints, *Hukou* constraints, and the age constraints for the mothers:

- a. Whether the community allowed all women to have two children.
- b. Whether the community allowed women to have two children if the first child was a girl, called the ‘girl-exception’.
- c. Whether the community allowed women to have two children if the parents worked in special occupations, called the ‘occupation-exception’.
- d. Whether the community let women have two children if both parents are only-children, or one parent is only-child, or one parent and both grand-parents are only-children, called the ‘only-child-exception’.

For example, if a woman had a daughter and lived somewhere with the girl-exception and the birth gap constraint was four years, then the woman is defined to be eligible for two children when her daughter reached the age four. If she became eligible in 1992, then her status will be ‘yes’ from the 1993 survey wave onwards. We then define the local OCP strength as the proportion of women in each community who were eligible to have two children according to the above rules, denominating by the number of ever-married women aged from 20 to 49. Higher values of this measure indicate more relaxed local regulations, and this should be positively related to the number of siblings. Moreover, by controlling for

⁶ Marriage is the traditional and legal pre-condition of child bearing in China. The legal marriage age is 20 for women and 22 for men from 1980 onwards, and was 18 for women and 20 for men between 1950 and 1980. Children born with either parent under the age limit will be considered an illegal birth.

the local development index, and for nutrient intakes of adults, the potential correlation between parental preferences for child quality and the relaxation of the OCP is controlled for, so the variation in the measure of OCP strength should have no direct impact on child nutrient intakes except via its impact on fertility choice (Liu 2014).

Our OCP variable based on four types of exceptions is more comprehensive than some measures in the literature that relate women's eligibility to a limited number of exceptions (for example, Short *et al.* 2001 and Yang 2007), and this should increase the correlation between our measure of OCP strength and family size. Also, by defining this measure at the community level instead of considering the eligibility of individuals, this measure is less likely to affect the children's nutrient intakes through channels other than its influence on the family size.

For the FE approach, we estimate impacts from the change in sibling numbers, using the panel structure of the data, and also test other possible determine factors of children's nutrient intakes. In this approach, the dependent and independent variables are all measures of changes. The variables in the control vectors X and Z in equation (1) that can change over time are kept for the panel analysis, and those that are time-invariant, including the gender, whether being the first child in the family and the province of residence, are dropped from the vectors of control variables. In both the FE and IV analysis, to allow the effects on the nutrient intakes to interact with gender, we run regressions separately for boys and girls.

4. DATA AND DESCRIPTIVE ANALYSIS

The main source of data for the analysis is the China Health and Nutrition Survey (CHNS), provided by Carolina Population Centre and the Chinese Centre for Disease Control and Prevention. The CHNS survey is an ongoing international collaborative project, which amongst other things, aims to examine the effects of health, nutrition, and family planning policies and programs. It is one of the few surveys to provide information on variation in the local stringency of the OCP. The CHNS is a longitudinal survey, following the same households over time and collecting relationship information on all family members, even those outside the household. For example, all children of each woman are covered no matter whether living with the mother at the time of the survey. This helps us better define eligibility for having two children because the birth rules are based on all children that each woman ever had. The surveys cover economic status, such as household income and subsidies; social status such as *Hukou* and family composition; and parental investments in children such as the spending on inputs into their human capital formation. In particular, it has detailed data on individual diets, which provides our outcome measures.

4.1 Measures of the One-Child Policy Strength

Two sources are used to define the OCP regulations in each community. One source is the CHNS community level data, where the local exceptions for a second or third child are

reported by the public officials responsible for the OCP implementation. Three out of four exceptions covered by the questionnaire are used in defining the local OCP strength, in terms of whether all women in the community were allowed to have two children, the girl-exception and the only-child-exception.⁷ The second source is government documents about provincial OCP regulations, which are used to identify the girl-exception, the only-child-exception and the occupation-exception when the CHNS information is missing. In addition, these documents provide information about the constraints on applying the exceptions in each province in different years, including the birth-gap for consecutive children, the age limitations on mothers and the *Hukou* requirement.⁸ The four exceptions, along with the constraints, are then applied to the families in the sample to define women's eligibility for having two children.

4.2 Nutrient Intakes

The CHNS has respondents keep three-day records of their food consumption, in terms of meals per person per day, and daily food intake. This information is then cleaned and reported as three-day average daily intakes for fat, protein, energy and carbohydrate, where energy is measured in kilocalories (kcal) and the others are in grams. Among these, daily energy is typically used as a general measure of the nutrition level. Diets with more fat and protein are traditionally seen as better nutrition than carbohydrate-rich diets in China (Zhang *et al.* 2016) and the higher income elasticity for protein and fat reflects this. Thus, intakes of these two nutrients should capture dietary preferences and reflect investments in each child.

We use 'relative nutrient intake' as the dependent variables in the regressions. This is the ratio of each child's three-day average nutrient intake level to the Recommended Nutrient Intake (RNI) for their age-gender group. The RNI are from the Chinese Dietary Reference Intakes (DRIs), which are based on Recommended Dietary Allowances (RDA), and are the level of intake that satisfies most individuals' needs in certain age, gender, and physiological status groups, and is set as $1.2 \times \text{EAR}$ (Estimated Average Requirement).⁹ For infants, we use

⁷ The last exception was to allow a second child if the first child was handicapped. We exclude it in this paper because we are not able to tell whether the first child was handicapped for each family, and hence cannot apply this exception to the families.

⁸ For example, Liaoning required the mother to have rural *Hukou* when applying the girl-exception, and further required the first girl to have rural *Hukou* since 2003; in 1990, Hubei set the birth-gap constraints to be four years and required mothers to be over 28 when having the second child, and relaxed the age limit of the mothers to be 25 in 1999 and changed the birth-gap constraint to be two years and only applied to women aged below 28.

⁹ When the standard deviations (SD) for the group are available and the nutrient intakes of the population are normally distributed, RNI equals EAR plus 2SD, otherwise it is set as $1.2 \times \text{EA}$. In this research, $1.2 \times \text{EA}$ is used as the reference level for each nutrient due to the lack of information on the SD of the population. RNI for boys and girls aged under 18 are exactly the same for fat but differ in some age ranges for protein. For energy, the RNI is higher for boys than for girls at all ages.

Adequate Intakes (AI) as the reference nutrient intake level instead of RNI, due to lack of information. The AI also suggests the level of nutrients that should satisfy most individuals' daily needs although not as accurately as RNI. Thus, we consider physical differences between boys and girls, and between children of different age groups, which makes inter-personal comparison of nutrient intakes easier. Ratios smaller than one show insufficiency of the nutrients and ratios exceeding one indicate that the daily intakes are more than sufficient.

4.3 Sample Selection

Seven waves of CHNS data, from 1991, 1993, 1997, 2000, 2004, 2006 and 2009 are used. Nine provinces are covered by the survey, except Liaoning was absent in 1997 and Heilongjiang was absent in 1991 and 1993. We do not require a balanced sample for our analysis so we begin by using all the observations. We include children with biological, step and foster parent-child relations and prioritize child-parent dyads where the parent is in the same household as the child, since that is the parent who should determine the diet of the child. We only consider children of the majority Han ethnicity, and exclude Han children living in communities where Han are not a majority, since diets in non-Han areas may be quite different.

We also exclude households whose residents include children aged over 18, since diets of these adult children are less likely to be determined by their parents. In addition, children born before 1976 are excluded because the fertility choices of those families were not affected by the OCP (Qian 2009 and Zhang *et al.* 2016). Households with twins are also excluded from our sample, since the investments by parents in the twins and in the siblings of the twins may not be comparable with those for other children (Li *et al.* 2008, Rosenzweig and Zhang 2009 and Liu 2014). We further require that children are present in at least two waves of the survey, since we need this for the fixed effects analysis. In order to calculate community level OCP strength, we also drop communities with fewer than three ever-married women aged from 20 to 49.¹⁰ As a result of the selection criteria, our sample is not a balanced sample although it is still quite large with almost 11,000 observations. The number of communities also varies over time, since communities without eligible children are excluded from the sample. For rural communities the range is from $n=105$ in 2009 to $n=126$ in 2000, and from $n=46$ in 2009 to $n=69$ in 2000 for urban communities.

4.4 Descriptive Analysis

The mean values of the outcome variables, instrumental variable, and control variables are reported in Table 1. The mean fat intake is more than recommended, while energy and protein intakes are, on average, below recommendations. These measures already control for gender and age effects and so the higher intakes of boys suggest some son preference.

¹⁰ Women who are reported to be 'married', 'divorced', 'widowed' or 'separated' at the time of the survey are treated as 'ever-married',

Table 1: Descriptive Statistics for the Estimation Sample, CHNS, 1991-2009

	All Children	Boys	Girls
<i>Outcomes</i>			
Relative Energy Intakes	0.77	0.78	0.76
Relative Protein Intakes	0.69	0.71	0.67
Relative Fat Intakes	1.55	1.59	1.50
<i>Instrument</i>			
Community OCP Strength ¹	0.82	0.83	0.82
<i>Children's Characteristics</i>			
Number of Siblings	0.85	0.79	0.94
Age (2-digit)	9.52	9.45	9.61
First-born Child	63.19%	60.71%	66.12%
Urban <i>Hukou</i>	30.89%	30.43%	31.43%
<i>Parents' Characteristics</i>			
Mother's Years of Schooling	7.16	7.18	7.14
Father's Years of Schooling	8.62	8.61	8.63
<i>Community Characteristics</i>			
Median Per-capita Real Household Income (yuan) ²	3856	3864	3847
Local Development Index ³	52.12	51.78	52.52
Median Energy Intakes of Adults (kcal) ⁴	2377	2574	2208
Median Protein Intakes of Adults (gram)	69	75	64
Median Fat Intakes of Adults (gram)	63	67	59
<i>Province of Residence</i>			
Liaoning	6.25%	5.82%	6.76%
Heilongjiang	8.05%	7.92%	8.21%
Jiangsu	12.28%	12.64%	11.85%
Shandong	11.00%	10.73%	11.33%
Henan	13.52%	13.43%	13.64%
Hubei	15.74%	15.21%	16.36%
Hunan	12.28%	12.20%	12.37%
Guangxi	15.65%	16.73%	14.37%
Guizhou	5.22%	5.32%	5.11%
<i>Survey Year</i>			
1991	19.42%	19.25%	19.62%
1993	21.12%	20.91%	21.37%
1997	18.76%	18.51%	19.05%
2000	16.90%	17.02%	16.76%
2004	10.38%	10.41%	10.36%
2006	8.52%	8.71%	8.29%
2009	4.90%	5.19%	4.55%
Total Number of Observations	10,988	5,958	5,030

Notes:

Values in the table are means unless noted otherwise.

1. Community OCP strength is the proportion of women eligible to have two children according to the local rules, with the number of ever-married women aged 20 to 49 in each community as the denominator.
2. Median per-capita household income is adjusted to 2009 prices.
3. The index is provided in the CHNS database and is constructed from community level indicators of: population density, education, sanitation, housing, transportation, communications, health, market development, economic development, diversity, and indicators of modern markets and social services.
4. The median intake of each nutrient item is calculated based on the raw intake levels for adults in the same community in each wave. In each community, the base population is all adults aged above 18 for the column 'All Children', all males aged above 18 for the column 'Boys' and all females aged above 18 for the column 'Girls'.

Relatedly, the number of siblings has a mean of almost one for girls and just 0.79 for boys while girls are more likely to be first-born than boys. It seems that parents whose first-born is a girl are more likely to have another child, which is a result of the girl-exception to the OCP and is also indicative of the son preference. The other Table 1 nutrient intake statistics are for adults; these are not normalized by recommended level, so the higher values for males than for females may reflect their stature and activity.

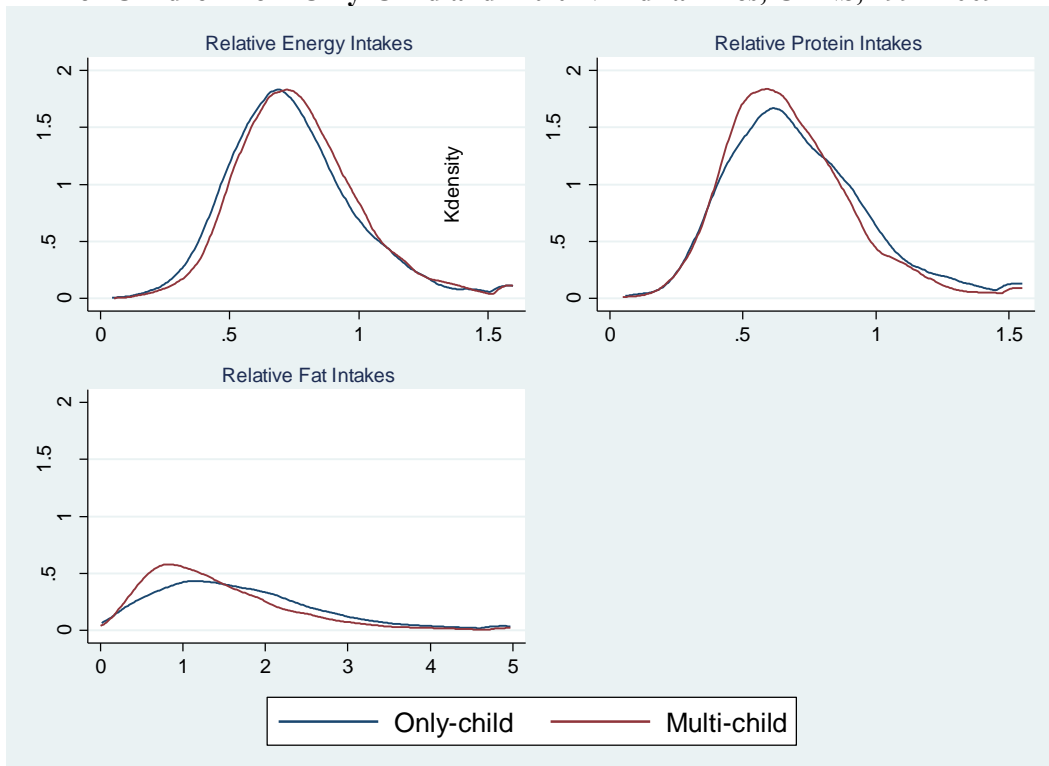
Figure 2 shows graphs of kernel density functions for the three nutrients, pooling all survey waves and comparing children from only-child families with others.¹¹ Children with no siblings had a greater density at the lower level of relative energy intakes than children with siblings, and greater density at the higher level of relative protein and fat intakes. Children from multi-child families had more density around the mean for protein intakes while only-children had heavier tails at the higher intake levels. This pattern is even stronger for relative fat intakes; children with siblings have more density in the range of 0.5 to 1.5 while only-children have a flatter distribution and much lower density at values below 1. Given preferences of parents and children for high protein and high fat food (Zhang *et al.* 2016), these patterns suggest that children from sole-child families get more dietary resources than children with siblings.

To show how the intakes vary over time, we plot the mean intakes for boys and girls in each wave for children from only-child families and multi-child families in Figure 3. The trend is for energy and protein intakes to go down, whereas fat intakes increased. Previous research using the CHNS data also notes that total energy intakes of Chinese children have decreased slightly over time while there has been an increase in dietary diversity and in fat intakes (Liu *et al.* 2013). The time trends for boys and girls had no obvious differences.

In terms of differences in intakes for children from multi-child families versus children from only-child families, these are most apparent for protein and fat intakes. Only-children had higher relative protein intakes than for those with siblings in all waves, irrespective of gender. For fat intakes, the gaps varied from almost none in the 2000 survey wave to larger gaps in the earlier and later waves. Interestingly, girls with siblings in the 2009 survey wave had higher fat intakes than only-child girls, reversing the pattern seen in all prior waves. Whether these raw patterns persist once we control for other characteristics of the child, of their family, and of the community – particularly the stringency of the local OCP regulations – is explored with the regression analyses in the next section.

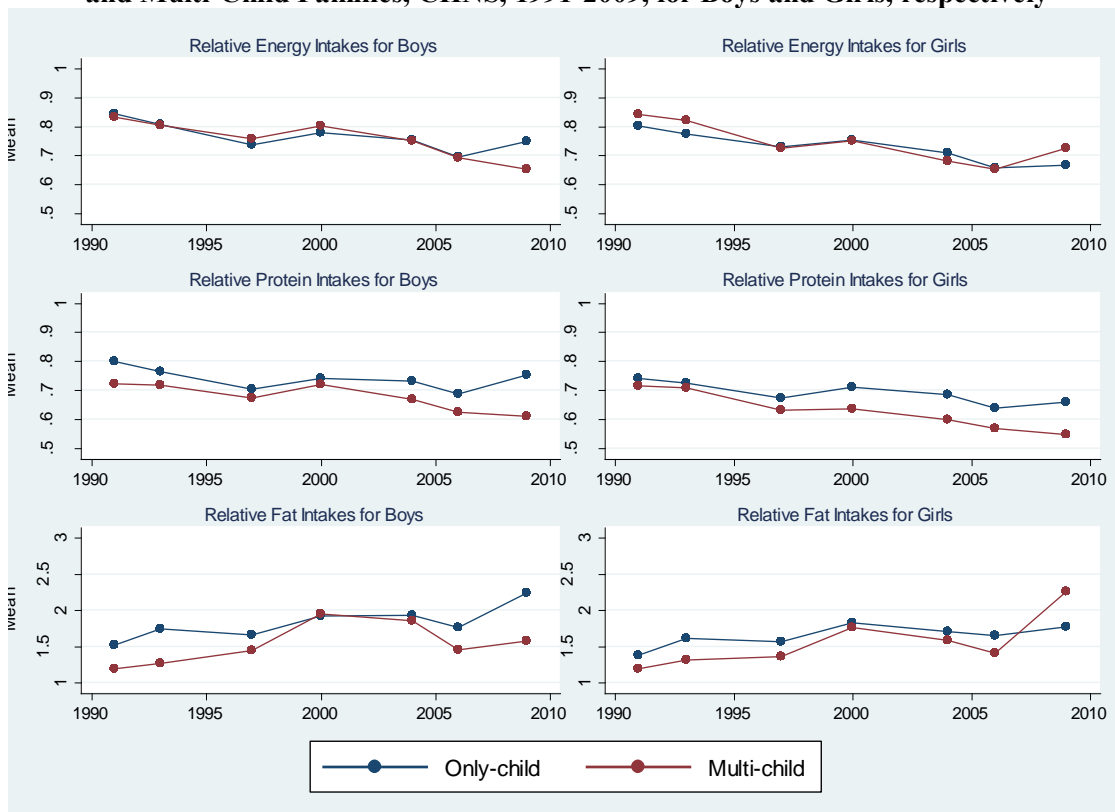
¹¹ For display purposes, relative intakes are trimmed at the 99th percentile. The raw distributions are all normal distributions with long tails, with the upper tails for fat being extremely long.

Figure 2: Kernel Densities Functions of the Relative Nutrient Intakes for Children from Only-Child and Multi-Child families, CHNS, 1991-2009



Notes: Sample from the CHNS, 1991-2009, Han children aged between 0-18 from Han communities. Relative intakes at the top one percent are grouped at the 99 percentile for each nutrient item.

Figure 3: Time Trends of the Relative Nutrient Intakes for Children from Only-Child and Multi-Child Families, CHNS, 1991-2009, for Boys and Girls, respectively



5. REGRESSION RESULTS

5.1 Impacts from Child Quantity

5.1.1. Cross-sectional Results

We begin by presenting the regression results from the OLS and 2SLS methods for the three measures of nutrient intakes, initially ignoring the panel structure of the data. For all estimation methods and outcomes studied, we present regressions for all children first, and then for boys and girls separately.

Table 2 has regression results from OLS and from both stages of the 2SLS approach for all children. The nutrient intakes are the dependent variables for the OLS regressions (columns 1, 4, 7) and for the second stage of the 2SLS (columns 3, 6, 9). The dependent variable for the first stage regressions (columns 2, 5 and 8) is the number of siblings. The effect of the instrumental variable on sibling numbers is positive and significant; as a higher share of women in a community were allowed to have a second child, under the local OCP exceptions, the number of siblings for each child increased. The F -tests for excluding the local OCP strength variable from the first stage are larger than 10, and so based on the rule-of-thumb from Bound *et al.* (1995), this indicates that there are no problems with weak instruments.

The OLS results suggest that nutrient intakes are lower for children with more siblings, although only for fat intakes are the effect statistically significant (at the 1% level). According to the OLS results, the addition of a sibling would be expected to reduce the fat intake of a child by about ten percent of the recommended level. The magnitudes of the trade-offs are much larger in the 2SLS results, ranging from -0.18 for dietary energy to -0.36 for fats, and these 2SLS results are statistically significant for energy (at the 10% level) and for protein (at the 5% level). An added variable form of the Durbin-Wu-Hausman (DWH) test for endogeneity that is due to Davidson and MacKinnon (1993) indicates that the difference between the OLS and 2SLS results for energy and protein are statistically significant, implying that the OLS results for these nutrients will be inconsistent due to the endogeneity of sibling numbers. That is not true, however, for fat intakes; the wide standard errors around the 2SLS results make them insignificantly different from the OLS results.

Thus, when boys and girls are pooled together, the valid models are 2SLS for energy and protein, and OLS for fats. The trade-off effects in these models range from -0.11 to -0.22 and are statistically significant at 10% (energy), 5% (protein), and 1% (fats) levels. So an additional sibling reduces nutrient intakes by between one-tenth and one-fifth of the recommended level, which is a substantial effect. This trade-off is more precisely measured for the more preferred nutrients of fats and proteins.

Table 2: Impacts on Relative Nutrient Intakes for All Children Aged under 18, CHNS, 1991-2009

	Relative Energy Intakes			Relative Protein Intakes			Relative Fat Intakes		
	OLS (1)	First Stage (2)	2SLS (3)	OLS (4)	First Stage (5)	2SLS (6)	OLS (7)	First Stage (8)	2SLS (9)
Instrument – Local OCP Strength		0.142** (0.027)			0.140** (0.027)			0.134** (0.027)	
Number of Siblings	-0.00239 (0.005)		-0.176+ (0.098)	-0.00017 (0.005)		-0.221* (0.100)	-0.107** (0.021)		-0.360 (0.381)
Female	-0.0199** (0.006)	0.192** (0.018)	0.0134 (0.019)	-0.0419** (0.005)	0.192** (0.018)	0.000495 (0.020)	-0.102** (0.030)	0.193** (0.018)	-0.0536 (0.068)
First-born Child	-0.00576 (0.008)	-0.707** (0.020)	-0.129+ (0.070)	0.00109 (0.006)	-0.708** (0.020)	-0.156* (0.072)	-0.069 (0.045)	-0.706** (0.020)	-0.249 (0.259)
Urban <i>Hukou</i>	0.0153* (0.007)	-0.186** (0.019)	-0.0192 (0.021)	0.0177* (0.007)	-0.192** (0.019)	-0.0275 (0.023)	0.0108 (0.036)	-0.190** (0.019)	-0.0403 (0.075)
Local Development Index	0.0614** (0.012)	-0.186** (0.036)	0.0251 (0.025)	0.0652** (0.011)	-0.203** (0.036)	0.0155 (0.026)	0.101* (0.047)	-0.171** (0.036)	0.0526 (0.099)
Total Number of Observations	10,988	10,988	10,988	10,988	10,988	10,988	10,988	10,988	10,988
Adjusted R-squared	0.137	0.528	0.029	0.169	0.528	-0.023	0.18	0.53	0.17
F-statistics for excluding local OCP strength		28.75			27.92			25.54	
P-value for added variable Hausman Test			0.059			0.013			0.509

Notes:

The dependent variable for the first stage is the number of siblings. All regressions control for age, age squared, years of schooling of parents, median household income, median nutrient intakes, and have dummy variables for province of residence and survey waves. Robust, clustered, standard errors, which allow for the potential correlation for the same children in different waves, are shown in parentheses.

** p<0.01, * p<0.05, + p<0.1

In the OLS results, but not in the 2SLS results, there is a statistically significant gender effect, with girls getting intakes (relative to requirements) of energy, protein, and fats that are 2%, 4%, and 10% lower than for boys, all else the same. Based on the DWH test results, only the OLS model for fats would be considered as insignificantly affected by endogeneity, and so the most valid inference is that some form of gender bias causes girls to have fat intakes that are one-tenth of the recommended level lower than for boys. The other (related) gender effect is in terms of sibling numbers, with a girl having 0.2 more siblings than a boy, on average. To further examine possible gender heterogeneity in the trade-off effect from sibling numbers on nutrient intakes, we run the OLS and 2SLS regressions separately for boys and girls, with results reported in Table 3.

The gender disaggregated results in Table 3 suggest that the trade-off effects for energy and protein intakes seen in the pooled sample are due to boys, where the DWH tests support the 2SLS results over the OLS results. In contrast, the trade-off between sibling numbers and the quality of diet observed through the lens of fat intakes is apparent for both boys and girls according to the valid OLS models. In fact, the trade-off appears to be twice as large for girls, with a coefficient of -0.15 compared to -0.07 for boys; in other words, an extra sibling would reduce the fat intake of a girl by about 15% of the recommended intake.

The other feature of the results in Table 3 is for the first-stage models in columns (2), (5) and (8). It appears that the number of siblings that a girl has is more strongly affected by external factors than is the case for boys. Thus, for girls the coefficients on the strength of the local OCP is about 0.15, versus 0.12 for boys, so as more women in a community are allowed to have a second child, if the existing child is a girl she is more likely to get an extra sibling compared with if the child is a boy. Similarly, while having urban *Hukou*, or living in an area with higher levels of the local development index both reduce the number of siblings that a girl has by about one-quarter, for boys the effect is only about one-half as large. Thus, for a girl the external constraints on fertility decisions of the family that may yield an extra sibling may be more important than for a boy, which is also seen in the first-born child indicator; a boy who is a first born has significantly fewer siblings than a first-born girl, which again reflects a form of son-preference.

In summary, the cross-sectional results suggest that fat intakes exhibit a significant trade-off between child quantity and quality for both boys and girls. The reduction in fat intakes for girls, due to having an extra sibling, is approximately twice as large as the reduction for boys. For energy and protein intakes, the significant trade-off seen in the 2SLS results is only for boys. While these cross-sectional results have plausible treatments for the number of siblings being an endogenous choice variable they do not deal with child-specific heterogeneity and so we now turn to panel regression results that look at within-child variation over time.

Table 3: Impacts on Relative Nutrient Intakes for Children Aged Under 18 by Sex, CHNS, 1991-2009

	Relative Energy Intakes			Relative Protein Intakes			Relative Fat Intakes		
	OLS (1)	First Stage (2)	IV (3)	OLS (4)	First Stage (5)	IV (6)	OLS (7)	First Stage (8)	IV (9)
<i>Boys (5958 observations)</i>									
Instrument - Local OCP Strength		0.126** (0.033)			0.125** (0.033)			0.122** (0.032)	
Number of Siblings	0.00604 (0.008)		-0.319+ (0.184)	0.00647 (0.008)		-0.357+ (0.195)	-0.0661** (0.026)		-0.396 (0.503)
First-born Child	0.000774 (0.009)	-0.759** (0.025)	-0.246+ (0.141)	0.00366 (0.010)	-0.759** (0.025)	-0.273+ (0.150)	-0.0198 (0.036)	-0.758** (0.025)	-0.27 (0.386)
Urban <i>Hukou</i>	0.0238* (0.010)	-0.117** (0.025)	-0.0179 (0.027)	0.0218* (0.011)	-0.121** (0.025)	-0.0261 (0.030)	0.0575 (0.038)	-0.117** (0.025)	0.0154 (0.080)
Local Development Index	0.0519** (0.015)	-0.141** (0.045)	-0.000539 (0.036)	0.0776** (0.016)	-0.150** (0.045)	0.016 (0.040)	0.133* (0.056)	-0.131** (0.045)	0.0835 (0.093)
Adjusted R-squared	0.14	0.557	-0.205	0.151	0.557	-0.215	0.328	0.558	0.307
F-statistics for excluding local OCP strength		14.89			14.73			14.05	
P-value for added variable Hausman Test			0.036			0.024			0.503
<i>Girls (5030 observations)</i>									
Instrument - Local OCP Strength		0.156** (0.041)			0.151** (0.041)			0.145** (0.042)	
Number of Siblings	-0.00865 (0.007)		-0.0678 (0.110)	-0.00476 (0.006)		-0.131 (0.097)	-0.149** (0.035)		-0.384 (0.576)
First-born Child	-0.0115 (0.014)	-0.647** (0.033)	-0.05 (0.070)	-0.00033 (0.008)	-0.647** (0.033)	-0.0824 (0.064)	-0.12 (0.090)	-0.646** (0.033)	-0.273 (0.337)
Urban <i>Hukou</i>	0.00595 (0.012)	-0.262** (0.029)	-0.0104 (0.030)	0.0142 (0.010)	-0.270** (0.029)	-0.0216 (0.029)	-0.0426 (0.063)	-0.270** (0.029)	-0.109 (0.141)
Local Development Index	0.0496** (0.018)	-0.247** (0.058)	0.0335 (0.040)	0.0421** (0.014)	-0.269** (0.058)	0.00512 (0.033)	0.109 (0.080)	-0.221** (0.058)	0.0523 (0.198)
Adjusted R-squared	0.124	0.506	0.111	0.189	0.505	0.095	0.1	0.508	0.093
F-statistics for excluding local OCP strength		14.41			13.39			12.14	
P-value for added variable Hausman Test			0.592			0.176			0.693

Notes: See Table 2.

5.1.2. Panel Results

The results of using child-specific fixed effects are reported in Table 4. The *F*-tests reported at the bottom of the table are all significant at 1% level, showing that the fixed effects cannot be excluded from any of the equations. In other words, there seem to be important, unobserved, child-specific effects that affect nutrient intakes and the omission of these from the previous models may have affected the results for tests of the quality-quantity trade-off.

The fixed effects results show that the negative effect of siblings on nutrient intakes is larger for girls than for boys, for all three nutrients, and is larger for fats than for the other two nutrients. An increase in the number of siblings (by one sibling) will decrease a child's fat intake by 13% of the recommended level for girls and by 11% for boys (and also by 13% in the gender-pooled sample). This is consistent with other studies using CHNS that find an only-child has a higher chance of being overweight (Zhang *et al.* 2016). For relative energy and protein intakes, the negative effects are smaller and statistically insignificant, in contrast to the cross-sectional 2SLS results.

5.2 Impacts from Other Factors

Although the main purpose of the regression models is to test for a quality-quantity trade-off that operates through the effect of sibling numbers on nutrient intakes, the effects of some other factors are also notable. In the OLS cross-sectional results, the local development level appears to have a significant positive effect on all three nutrient intakes, and although this effect disappears in the 2SLS results (presumably because local development correlates with the strength of OCP) it reappears for protein intakes in the fixed effects results.

There are similar changes in results for whether the child holds urban *Hukou*, which has positive effects on energy and protein intakes in the OLS results but negative effects, especially on protein, in the fixed effects results. Amongst the factors that are not measuring sibling numbers, the only consistently significant impact in all regression models, for all nutrient intakes and for both boys and girls, is the median nutrient intake level of the adults in the community. It seems that a significant factor in the children's nutrient intakes is what other people (adults) eat in the community.

Table 4: Panel Regression Results, CHNS, 1991-2009

	<u>Relative Energy Intakes</u>			<u>Relative Protein Intakes</u>			<u>Relative Fat Intakes</u>		
	All Kids (1)	Boys (2)	Girls (3)	All Kids (4)	Boys (5)	Girls (6)	All Kids (7)	Boys (8)	Girls (9)
Number of Siblings	-0.0169 (0.015)	-0.00884 (0.027)	-0.0134 (0.018)	-0.0169 (0.016)	-0.00193 (0.035)	-0.0158 (0.017)	-0.127* (0.050)	-0.108 (0.075)	-0.128+ (0.075)
Urban <i>Hukou</i>	-0.0175 (0.011)	-0.021 (0.016)	-0.0153 (0.016)	-0.0282* (0.012)	-0.0364* (0.017)	-0.0149 (0.015)	-0.0134 (0.041)	0.00666 (0.058)	-0.0148 (0.056)
Local Development Index	0.0275 (0.024)	0.0198 (0.035)	0.0381 (0.032)	0.0439+ (0.023)	0.0749* (0.037)	0.00713 (0.028)	0.046 (0.084)	-0.0323 (0.101)	0.239 (0.146)
Total Number of Observations	10,988	5,958	5,030	10,988	5,958	5,030	10,988	5,958	5,030
Total Number of Individuals	3,760	2,026	1,734	3,760	2,026	1,734	3,760	2,026	1,734
Adjusted R-squared	0.089	0.078	0.088	0.091	0.07	0.118	0.143	0.248	0.084
F-stat for all $u_i=0$	1.396**	1.277**	1.487**	1.27**	1.251**	1.215**	1.585**	1.216**	1.732**

Notes

All regressions control for age, age squared, years of schooling of parents, median household income, median nutrient intakes, and dummy variables for survey waves. Robust standard errors, which allow for the potential correlation for the same children in different waves, are shown in parentheses.

** $p < 0.01$, * $p < 0.05$, + $p < 0.1$

6. CONCLUSIONS AND DISCUSSION

In this paper we have tested for the existence of a trade-off between child quantity and quality in Chinese families. We used seven waves of CHNS data, with the quality of children measured by their intakes of energy, protein and fat, relative to recommended intakes for their age and gender. Two mechanisms are used to deal with possible bias in the estimation of causal effects coming from the endogeneity of the fertility choices. The first one is the 2SLS approach using the local strength of one-child policy restrictions as the instrumental variable when estimating sibling numbers (which are the quantity variable in the trade-off analysis). The second one is the fixed effects approach using the panel structure of the CHNS data to examine how changes over time in nutrient intakes of the same child are related to the changes in their sibling numbers. Together with the OLS results, this paper provides a comprehensive test of the quantity-quality trade-off theory in the context of China.

In general, the results are supportive of there being a quantity-quality trade-off with an extra sibling reducing intakes by between one-tenth and one-fifth of the recommended level. The negative effect of an extra sibling on fat intakes is somewhat larger for girls than for boys and is still apparent when child-specific fixed effects are used. In contrast, the effects on energy and protein intakes are more apparent for boys than for girls, and are not statistically significant when fixed effects are used. Since fat is a preferred nutrient in China (Zhang *et al.* 2016), even though current intakes are above recommended levels, the strength of the results for fat intakes, and the gender differences are consistent with son preferences.

Although there is reasonable consistency between the various estimation methods, a possible weakness of the IV method is that our instrumental variable does not perfectly identify the fertility constraints (despite the strength of our instrument in the first stage regression). As noted in other studies, the penalty for the unsanctioned birth can affect people's fertility choices (Liu 2014) and this is a variable that we did not include due to data limitation. Also, the reduced influence of OCP on fertility choices in later years may weaken its identification power. The fixed effects approach does not suffer from this potential weakening, and is useful in capturing the within-person changes to avoid omitted variable bias, but only a minority of the sampled children experiencing a change in sibling numbers during the duration of the survey.

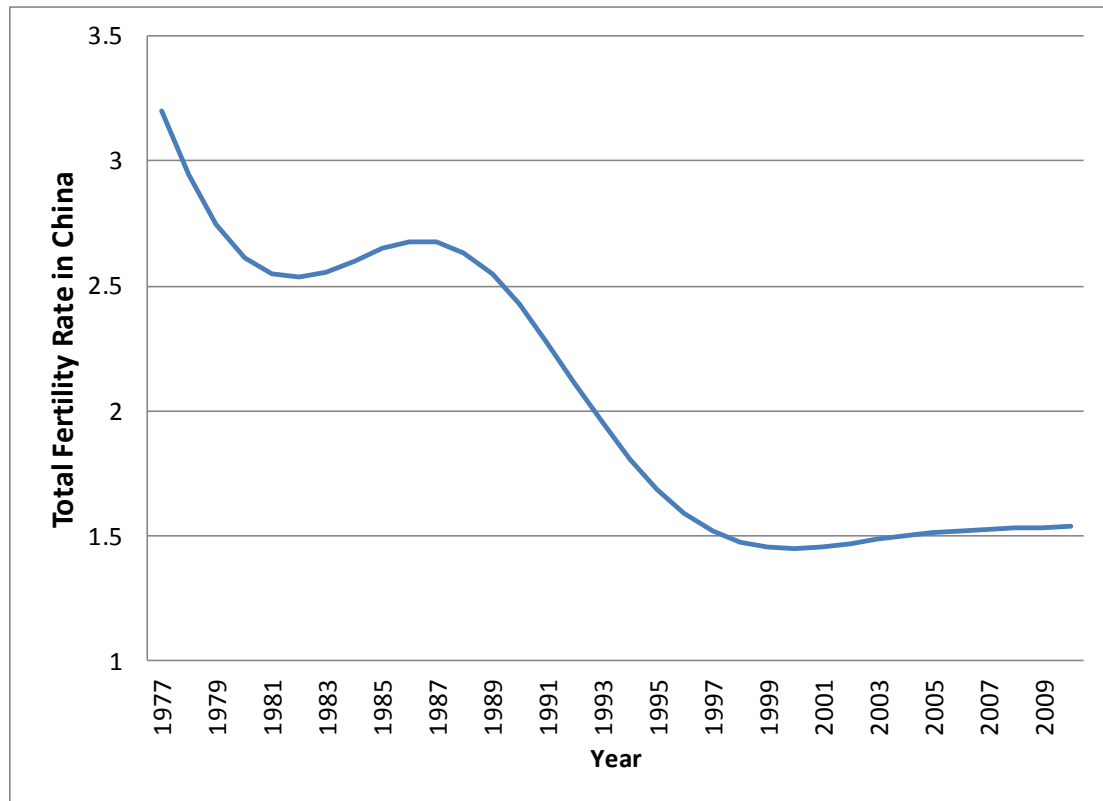
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Appendix 1
Time Trend of the Total Fertility Rate in China, 1977-2009



Sources

World Development Indicators

<http://data.worldbank.org/indicator/SP.DYN.TFRT.IN>, retrieved September 2016.

Source Note

Total fertility rate represents the number of children that would be born to a woman if she were to live to the end of her childbearing years and bear children in accordance with age-specific fertility rates of the specified year.

Appendix 2
Sample Distribution of Children by Urban and Rural Location and *Hukou* CHNS
1991-2009

	Boys		Girls		All Children	
	Rural	Urban	Rural	Urban	Rural	Urban
Rural <i>Hukou</i>	88.3%	13.8%	87.9%	11.1%	88.1%	12.5%
Urban <i>Hukou</i>	11.7%	86.2%	12.1%	88.9%	11.9%	87.5%
Number of Observations	4,460	1,502	3,765	1,266	8,225	2,768

Notes

Sample is from CHNS, 1991 to 2009 waves, all children aged below 18, Han ethnicity with valid information of nutrient intakes, personal characters and parents' characters and living in a community where Han is the major ethnicity.