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DO FERTILITY CONTROL POLICIES AFFECT HEALTH IN OLD AGE? EVIDENCE FROM CHINA'S ONE-CHILD EXPERIMENT[1]

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ABSTRACT

How do fertility control policies contribute to the welfare of women, and their husbands, particularly as they get older? We consider whether the reduction in fertility resulting from population control policies has had any effect on the health of elderly parents in China. In particular, we examine the influence of this fertility decline, experienced due to China's one-child policy, on several measures of the health of parents in middle and old age. Overall, our results suggest that having fewer children has a positive effect on self-reported parental health but generally no effect on other measures of health. The results also suggest that upstream financial transfers have a positive effect on several measures of parental health. Copyright © 2014 John Wiley & Sons, Ltd.

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1. INTRODUCTION

As Schultz (2008) has recently stressed, there have been very few studies of the long-run effects of policy-induced changes in fertility on the welfare of women or, we would add, their partners. Instead, most studies focus on the immediate or short-term effects of variations in fertility. For example, studies have examined the effects of the number of children one has on one's labour supply and wages, the stability of one's marriage and the educational outcomes of one's children (see, e.g. Haveman and Wolfe, 1995; Lundberg and Rose, 2002; Li *et al*., 2009). In addition, a number of studies have examined how the number of children one has affects parental health, particularly maternal health (Hurt *et al*., 2006). However, these studies have been restricted primarily to the period of the child's birth and the first few years of the child's life. There is a dearth of studies examining the long-term consequences for parental health of variations in fertility due to family planning policies, even though health is an accumulative outcome, as was emphasized by Strauss and Thomas (1998). Hence, the effects of children on parents' health during childbirth and when the children are young are likely to persist into old age.

We examine the consequences of China's one-child policy for the health of elderly parents. To the best of our knowledge, this is the first study in the economics literature to examine the long-term health effects of shocks in fertility resulting from population control policies. The issue is especially important in the context

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of China, as it has been confronting a rapidly ageing population. In 2000, the proportion of people of working age (15–59 years) was one of the highest in the world. However, by 2015, a smaller number of people will be joining the labour force, whereas a large number will be retiring. As a consequence, the size of the working age population will decline rapidly (UNFPA, 2006). China's old age social security system is still embryonic, meaning that the elderly parents in China are largely dependent on their adult children for support. For people aged 85 years and above, almost 80% are dependent on their children or other relatives for financial support. Around 70% of people aged 60 years and above live with their children or other relatives, and just 0.8% live in institutions (UNFPA, 2006). China's one-child policy has already substantially reduced the numbers of adult children who are able to look after their ageing parents, and this trend is set to increase in the future. The emerging crisis is dramatized by the oft-repeated observation in China that when people currently in the workforce retire, one couple will have to support four parents—the '1-2-4' phenomenon: one child, two parents and four grandparents (World Bank, 1997).

There are clear fertility trade-offs for old age health. On one hand, having more children could have a negative effect on health due to the physical and psychological stresses on parents when the children are young (and maybe even when they are older). There are studies in the medicine and psychology literature suggesting that having more children has an adverse impact on a range of parental health conditions, including mental health (see, e.g. Gove and Geerken, 1977), weight gain and obesity (see, e.g. Weng *et al*., 2004), cardiovascular disease (see, e.g. Zhang *et al*., 2009) and breast and cervical cancers (Ness *et al*., 1994). Many of these health outcomes, such as weight gain following the birth of children, have been shown to affect fathers as well as mothers. There is also a considerable amount of evidence that these health conditions can either persist into old age, as is the case for weight gain, or only become evident in old age, as for cardiovascular disease and stroke. Moreover, having more children also imposes financial constraints, because parents cannot borrow against the future earnings of their children when the latter are young. This means that more children could reduce the parental consumption of things that contribute to future health, such as good nutrition or health care. On the other hand, having more children could be beneficial for health in old age, as it increases the amount of care and financial support that elderly parents will receive from their adult children. There is also probably psychological value in having adult children when one is an elderly parent. Our main estimates of the instrumented effect of fertility on parental health in old age reflect the net result of these positive and negative effects. In addition, we also examine the co-habitation of adult children and financial transfers, to provide evidence of some of the mechanisms driving the net effect.

We use a number of health measures, as the relationships between various socio-economic determinants of health outcomes have been shown to be sensitive to different measures of health (Frijters and Ulker, 2008). Specifically, we consider five health measures in our attempt to capture the effects of the number of children on health outcomes. These are self-reported health, difficulties in performing activities of daily living (ADL), mental health, blood pressure and BMI. The empirical findings suggest that having more children has an adverse impact on the self-reported health of elderly parents. However, in general, we do not find the number of children to have any impact on other measures of parental health. Our results indicate that the effects are generally the same for maternal and paternal health. The findings suggest that upstream financial transfers have a positive effect on parental health. Our results have important policy implications for China in particular and developing countries in general. Of particular relevance to China is the fact that our results indicate that the one-child policy, which reduced the number of children in the post-1979 period, may not place parents in worse circumstances in terms of health outcomes in their old age.

2. HOW DOES ONE'S FERTILITY AFFECT ONE'S HEALTH IN OLD AGE?

There is much epidemiological literature to suggest that child bearing has long-term health effects. As we consider five different measures of health—self-reported health, ADL, mental health, blood pressure and BMI—it is possible that the number of children one has will be related to these different health measures in different ways.

2.1. Number of children, self-reported health and activities of daily living

The expected relationship between the number of children and self-reported health at older ages is ambiguous. There is medical evidence that suggests there is an association between high parity and several chronic conditions and diseases later in life, such as cervical cancer, cardiovascular disease and diabetes (Kington *et al*., 1997). The reason for this increased risk of disease later in life may be partly biological, being associated with the metabolic or physiological changes associated with pregnancy (which affect women) but is related primarily to the chronic stress and adverse lifestyle factors associated with child rearing, which increase the risk of a stroke (which affect both men and women).

An association between high parity and better self-reported health is also plausible for several reasons. First, there is medical evidence of a protective effect of higher parity for several important cancers later in life. Specifically, cancers of the breast, endometrium and ovaries are much less prevalent later in life among women who have had many children (see, e.g. Kvale *et al*., 1994). Second, multiple pregnancies may increase contact with the medical system, leading to a better access to medical care as one ages, through learned health care-seeking behaviours (Kington *et al*., 1997). Third, children may support their parents in their old age, with more children equating to more support, which might affect their parents' self-reported health (McGarry and Schoeni, 1997).

Similar channels linking the number of children to self-reported health are likely to apply to the link between the number of children one has and one's capacity to engage in ADL. People who have higher (lower) levels of self-reported health are likely to be more (less) able to perform routine activities. One potential point of difference is that, with ADL, the positive correlation between high parity and ADL as a result of the support structures associated with adult children is likely to be stronger. The question then is whether the potential negative health effects that would impair ADL in old age as a result of having more children will outweigh the positive effects on ADL due to having better support structures.

2.2. Number of children and mental health

Higher parity is expected to be associated with better mental health later in life. Social isolation and low levels of social support are important risk factors for depression (Cappeliez and Flynn, 1993), and adult children help to reduce feelings of social isolation among elderly parents. Although childless persons may also have social networks, consisting of relatives, friends and neighbours, research suggests that these networks are less likely to have the long-term commitment that adult children do (Zhang and Hayward, 2001). There might also be a psychological aspect of having children. Glenn and McLanahan (1981, p. 421) suggested that 'the mere knowledge that offspring are there to help if needed could contribute substantially to the psychological well-being of the elderly'.

2.3. Number of children, body mass index and blood pressure

Several studies have found that having children is a major cause of obesity among women (see the studies reviewed by Weng *et al*., 2004). The mechanisms proposed to explain the association between the number of children and obesity in women include the metabolic changes associated with pregnancy and the physiological changes associated with accommodating living with small children, such as changes in diet and physical exercise. A number of studies have found that having children increases the prevalence of obesity among fathers as well as mothers (see, e.g. Kritz-Silverstein *et al*., 1997). The explanation for this finding is that the behavioural changes that are associated with raising young children affect fathers as well as mothers and that couples with many children may have less opportunity, in a time allocation sense, to focus on resuming the sorts of health behaviours that promote weight loss. There are many evidence that weight gain among mothers and fathers when raising young children persists into old age (see, e.g. Heliovaara and Aromaa, 1981). Similar mechanisms explain the relationship between the number of children and blood pressure, as the number of children and weight gain are highly correlated with high blood pressure and obesity. Both higher blood pressure and BMI are positively correlated with cardiovascular disease, diabetes and the risk of a stroke in old age (Cohen *et al*., 2006).

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3. DATA

We use pilot data from the China Health and Retirement Longitudinal Survey (CHARLS), the first publicly available data set on the elderly parents in China (Zhao *et al*., 2009; Smith *et al*., 2012). CHARLS includes data collected from 1563 randomly selected households from two provinces, Zhejiang and Gansu, between July and September 2008. CHARLS sampled one person aged 45 years and over, plus their spouse if one exists, in each household with an age-eligible person. Hence, CHARLS contains information on 2951 individuals aged 45 and older. Gansu and Zhejiang were chosen as the two provinces because one is the representative of a poorer inland province and the other is representative of a prosperous coastal province. Gansu, in the northwest of China, is one of the poorest provinces in China and has a large rural population. In contrast, Zhejiang, on the south-eastern coast, is one of the leading centres of Chinese industrialization and is far more urbanized than Gansu province, with a much higher proportion of exports. Incomes in Zhejiang have risen more rapidly than the Chinese average. Full details of the sampling method are provided on the CHARLS website.

Table I contains descriptive statistics for the number of children and other control variables used in the regression that potentially explain the health status of the respondents. Most of the control variables that we employ are standard in the literature examining the determinants of health status (see, e.g. Case *et al*., 2005). Specifically, we control for age, gender, education, marital status, place of residence (urban/rural), number of children and number of siblings. Yamashita (2008) finds that sibship size predicts some health outcomes in later life among a sample of the Mexican elderly parents. We use household consumption expenditure as a control variable rather than household income, as the former is a better proxy for permanent income and in particular is a better proxy in our context because our sample includes many households that are retired and do not have any current income. We control for total household wealth, because one's access to health care as one ages may be affected by the level of one's assets. Equally importantly, we also control for respondents' health status during childhood using self-reported health status during childhood. A respondent's health status during childhood can also be indicative of their living conditions, with implications for their health later in life (Case *et al*., 2005).

When examining the health status of the elderly parents in China, the CHARLS data have the advantage that it contains a rich set of variables on health status that would otherwise be difficult to obtain. We use five

Table I. Descriptive statistics

	Zhejiang		Gansu	
Major socio-economic variables	Mean	SD	Mean	SD
Number of children	2.585	(1.412)	3.935	(2.073)
Number of siblings	3.648	(1.829)	3.935	(2.073)
Age of household head	59.28	(10.71)	58.08	(9.943)
Gender (male = 1)	0.534	—	0.521	—
Marital status (married = 1)	0.857	—	0.835	—
Hukou status (rural = 1)	0.808	—	0.813	—
Years of schooling of household head	3.395	(3.410)	3.386	(3.925)
Health during childhood	2.207	(0.932)	2.272	(1.118)
Maximum years of schooling of any member of the household	9.790	(3.198)	9.299	(3.643)
Annual per-capita household consumption expenditure (Yuan)	8504	(7121)	5234	(5794)
Household wealth ('000 s Yuan)	111.4	(246.4)	18.71	(66.87)
Outcome variable				
Self-reported health status (1–5)	3.221	(1.405)	3.472	1.641
ADL (0–1)	0.0444	(0.1017)	0.1167	(0.1452)
BMI (0/1)	0.4866	(0.5001)	0.4490	(0.4977)
Mental health (0–1)	0.2137	(0.1738)	0.2682	(0.2435)
Blood pressure (0/1)	0.3362	(0.4726)	0.3681	(0.4826)
Number of observations	1425		1260	

ADL, activities of daily living.

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alternative, albeit related, measures of health status, similar to the approach adopted by Frijters and Ulker (2008). As our first measure, survey respondents are asked to rate their current health status on the Likert scale where 1 = excellent health (2.56%), 2 = very good health (11.18%), 3 = good health (20.55%), 4 = fair health (40.23%) and 5 = poor health (25.47%). Our second measure of health status is ADL, which is a composite index of the level of difficulty that the respondent has in performing a number of fairly normal and routine day-to-day activities or tasks. Specifically, we consider the degree of difficulty experienced by the respondent in performing the following 21 tasks: dressing, bathing, eating, getting in or out of bed, walking 100 m, walking 1 km, sitting for 2 hours, getting up from a chair, climbing several flights of stairs, stopping, kneeling or crouching, lifting 10 *jin* (equivalent to a heavy bag of groceries), extending one's arm, pushing or pulling large objects, urinating, doing household chores, preparing hot meals, shopping for groceries, managing money, making phone calls and taking medicine. Respondents' answers were coded as follows: 1 = 'I do not have any difficulty in performing the task'; 2 = 'I can perform the task, but only with difficulty'; and 3 = 'I cannot perform the task'. Our measure of ADL is the sum of the responses for all 21 tasks. The responses are normalized so that the maximum value of the composite ADL variable is 1 and the minimum value is 0. The mean value of ADL is 0.078, with a standard deviation of 0.129.

Similarly, our mental health variable is a composite measure where the respondents were asked about whether they had experienced a range of mental health symptoms in the past week. Specifically, respondents were asked if they felt bothered by things that did not usually bother them, if they had trouble keeping their mind on what they were doing, if they felt depressed, if everything that they did in their life was an effort, if they felt fearful, if their sleep was restless, if they were happy, if they felt lonely, if they felt sad, and if they could not get going. The responses were coded as 1 = 'rarely or none of the time'; 2 = 'sometimes'; 3 = 'a moderate amount of the time'; and 4 = 'most or all of the time'. This variable is also normalized so that it takes a value between 0 and 1, with a higher value indicating poorer mental health. The mean value of the mental health variable is 0.239, with a standard deviation of 0.211. We also use the BMI and blood pressure. Less than 5% of the people in the sample are underweight. Hence, we merged them into the 'normal' weight group and simply use whether a person is of normal weight or overweight, where overweight is defined as having a BMI greater than 25. This also allows us to estimate a simple probit model in which the BMI variable takes a value of 1 if overweight and 0 if normal/underweight. To measure blood pressure, the blood pressure of the individual was taken three times, and we take the average of these blood pressure measures to determine whether an individual has high or normal blood pressure. We convert the blood pressure variable into a binary variable, which is equal to 1 if a person has high blood pressure (above 140/80) and 0 otherwise.

In Figure 1, we plot the health status of parents against their number of children. We find that self-reported health status does not vary much with the number of children (Figure 1A). For people with no children, self-reported health is very high, indicating a worse subjective health condition. Figure 1B shows that mental health improves with the number of children. Figure 1C shows that ADL goes up as the number of children increases after one child. For parents with one child, however, ADL is smaller than having no children or two or more children. This indicates that the functional status of a person is lowered with the number of children beyond one, and it is also lower for those with no children. Figure 1D shows that the BMI is almost the same for parents with children. Slightly less than half of the parents have BMI in the normal category. The BMI for childless individuals and parents with eight or more children are very low, but this may not be a reliable result, as there are very few observations in these categories. Figure 1E shows that the proportion of parents with high blood pressure increases as the number of children in a family increases beyond one. It also shows that more than one-third of people with no children have high blood pressure.

4. EMPIRICAL STRATEGY

We estimate the following health equation

$$Health_i = \beta_0 + \beta_1 Child_i + \beta_2 X_i + \upsilon_i, \tag{1}$$

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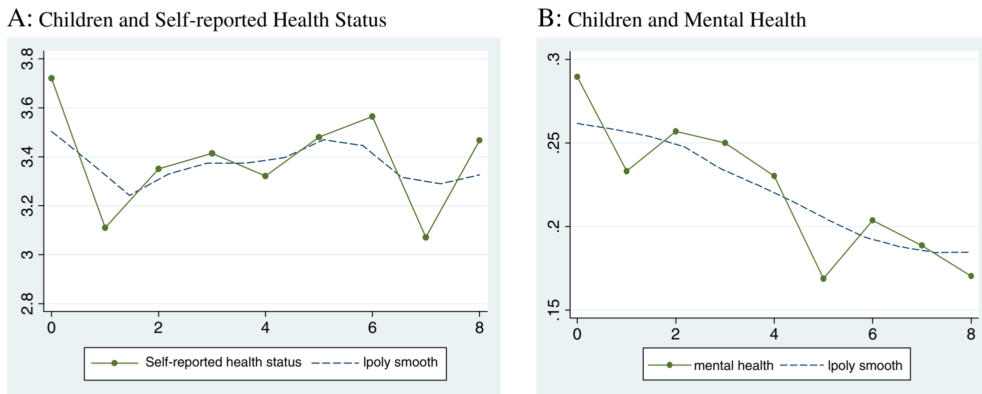

NB:The dashed line in each figure is fitted using a kernel-weighted local polynomial regression of the respective health status variable on the number of children.

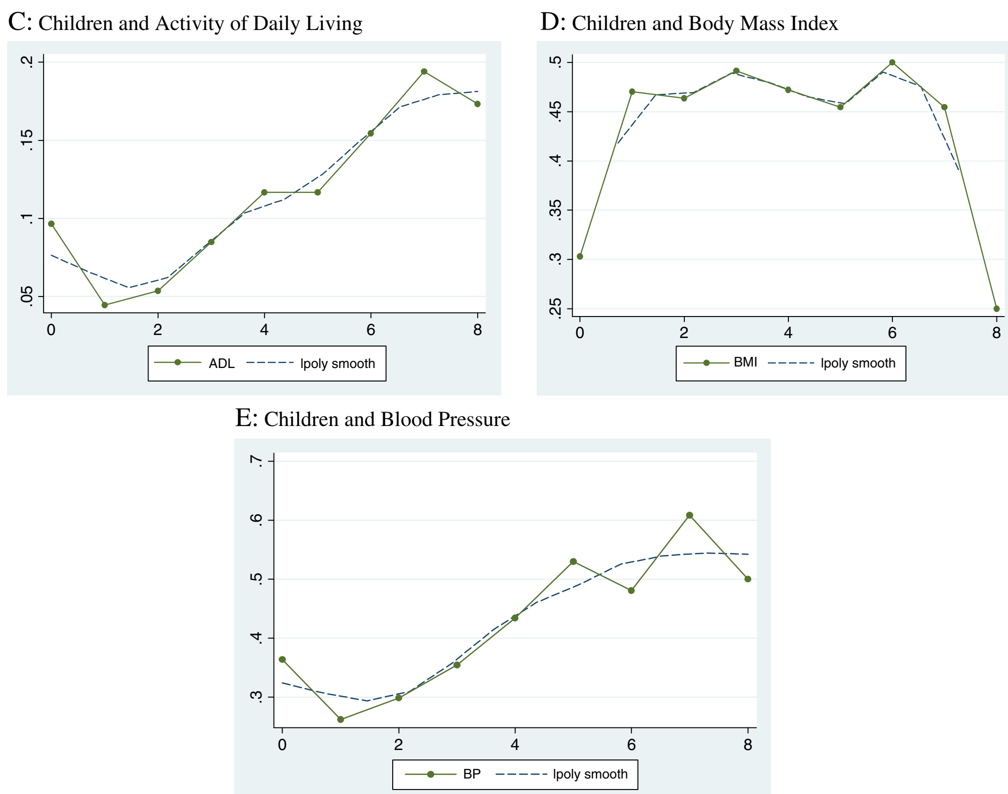

Figure 1. Numbers of children and the health status of elderly parents: A) children and self-reported health status; B) children and mental health; C) children and activity of daily living; D) children and body mass index; and E) children and blood pressure

where $Health_i$ is the health status of respondent i (aged 45 years or older), $Child_i$ is the total number of children born to individual i, and X_i is a vector of individual characteristics (age, gender, *hukou* status, marital status, education, health status during childhood, consumption, wealth, number of siblings and location of residence).

Estimating Equation 1 using ordinary least squares (OLS) is likely to yield biassed results, because the unobserved attributes that are correlated with the household's fertility decisions are likely to be correlated with the health status of the individual. The fertility level is chosen endogenously by parents and hence may be

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related to other unobservable parental characteristics that affect health outcomes. One method of addressing the endogeneity of the fertility decision is to use exogenous variation in the numbers of children that is generated by the natural occurrence of twins to isolate the causal effect of family size (see, e.g. Li *et al.*, 2008; Rosenzweig and Zhang, 2009). However, using twins to instrument for the number of children requires either a dedicated twins dataset (such as the Chinese Child Twins Survey) or a very large dataset, as the occurrence of twins in the population is very rare. CHARLS does not contain dedicated data on twins; hence, we were not able to use twins as an instrument for fertility.

Our instrument is based on the exogenous variation in family size caused by China's one-child policy. Specifically, our instrument is based on the 'one-son-two-child' rule, which allowed rural couples to have a second child if the first child was a girl. The one-child policy applied to all individuals of Han ethnicity, who constitute 92% of the Chinese population, and banned second births except in exceptional circumstances. Later, the one-child policy was partially relaxed to allow for regional variations in family planning policies and the 'one-son-two-child' rule in some areas allowed rural couples to have a second child if the first was a girl. Most of the respondents in the CHARLS dataset are Han Chinese. Hence, there is little variation in the ethnicity of respondents. In Zhejiang, more than 99% of residents are Han Chinese. Gansu has Tibetan minority counties, but Gansu's nine Tibetan minority counties (eight rural counties and one urban district) were excluded from the survey due to concerns about political sensitivity and the problems associated with language barriers.

The first-stage estimation is given by

$$Child_{ij} = \varphi_0 + \varphi_1 Rural_j \times Girl_i \times Policy_i + \varphi_2 X_{ij} + \xi_{ij} \tag{2}$$

where $Rural_j$ is a dummy variable =1 if a child was born in a rural area and $Girl_i$ is a dummy variable indicating if the first child was a daughter. $Policy_t$ is three different dummy variables corresponding to whether the child was born before 1973, 1976 or 1979.

We use three different instrumental variables (IVs) *(Girl*Rural*73, Girl*Rural*76* and *Girl*Rural*79)* separately in different specifications to reflect the fact that there were other population control policies, which were introduced in 1976 and 1973 apart from the one-child policy in 1979. Beginning in 1972, the policy 'Later [age], longer [the spacing of births], fewer [number of children]' was introduced as a precursor to the one-child policy. This earlier policy offered economic incentives to parents who spaced the birth of their children at least 4 years apart. Qian (2009) makes the point that most studies assume that the one-child policy only affected family size for cohorts born after 1979. However, as Qian (2009) argues, if the previous 4-year birth spacing law was enforced, the one-child policy should be binding for those born in 1976 and after. We also consider children born in or after 1973 to further allow for the effects of the 'Later [age], longer [the spacing of births], fewer [number of children]' on parental preferences with respect to birth spacing. If the child was born between 1973 and 1976, it is conceivable that parents were planning to have a second child when the one-child policy was introduced in 1979, as in the intervening period they were being encouraged to increase spacing between first and second births.

The first-stage results in Equation 2 were estimated using OLS. However, the health status variable in the second stage can be ordinal (self-reported health), dichotomous (BMI and blood pressure) or censored (ADL and mental health). Both the ADL and mental health variables have been normalized to be between zero and one, and many individuals reported that they had no problems. The second stage is estimated using either an ordered probit model, when the health variable is self-reported health, or a probit model, when the health variable is BMI or blood pressure. For ADL and mental health, we use a Tobit model, to account for the large number of zeros. As we use probit or Tobit models in the second stage, we do not apply traditional two-stage least-squares. Instead, we use reduced-form residuals in the second stage (Vella, 1993). We first obtain generalized residuals using OLS for the reduced-form first-stage equation then use the estimated residuals as an additional regressor in the second stage, which is equivalent to control function estimation.[2] We report robust standard errors; using bootstrapping with 500 replications gives similar results.

[2]The traditional example of the control function approach is Heckman's sample selection model that augments the outcome equation by an estimate of the Mills ratio.

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5. RESULTS

5.1. Results without instrumenting

We first present the results without allowing for endogeneity of fertility in Table II. Each specification in Table II uses covariates at the household and individual level as controls. The coefficient on the number of children is positive and significant when the dependent variable is self-reported health, ADL or BMI but insignificant when the dependent variable is blood pressure or mental health. These results suggest that respondents who have more children have statistically lower levels of self-reported health and more difficulty in performing daily living activities and are statistically more likely to be overweight *ceteris paribus*. However, the number of children is an endogenous variable, and therefore, these results might not reflect the causal effects of the number of children on parental health. We report results that address this issue in the succeeding texts.

5.2. Validity of instruments

Before we present our IV estimates, we check the reliability of our instruments. One important concern with the IV approach is the possible use of weak instruments, which tends to bias second-stage estimates and may weaken standard tests for endogeneity. We test whether the IVs are correlated with the endogenous regressor and orthogonal to the error process. We test the first condition by examining the fit of the first-stage regression, in which the number of children is regressed on *Girl*Rural*Policy*, together with control variables. The first-stage results are reported in Table III. The results for the one-child policy with time dummies indicate that the relaxation of the one-child policy will increase family size. The results of the *F*-tests for the three *Girl*Rural*Policy* variables are all in excess of 80.

The second condition to have a valid instrument is the requirement to satisfy the exclusion restriction. In our case, this means that the one-child policy only affects health through an increase in the family size. The exclusion restriction is not testable directly. We estimate a semi-reduced-form equation, in which fertility is instrumented, but the IV enters the second-stage regression directly (and naturally in the first-stage regression), to examine whether the instruments have an independent effect on parental health. The results do not indicate any significant effect of the one-child policy on parental health in any specification. Consequently, we conclude that both the first-stage and reduced-form results support the use of our instrument.

5.3. Instrumental variable results

The IV estimates in which health status is regressed on the number of children and the control variables are reported in Table IV. Overall, the results suggest that respondents who have more children have statistically lower levels of self-reported health at the 10% or lower when the number of children is instrumented using the *Girl*Rural*Policy* variables. However, for the other measures of health, the results suggest that the number

Table II. Effect of the number of children on parental health status (not instrumented)

	(1)	(2)	(3)	(4)	(5)
	Self-reported health	ADL	Mental health	BMI	Blood pressure
Number of children	0.064**	0.01**	−0.005	0.08**	0.030
	(0.019)	(0.00)	(0.005)	(0.03)	(0.026)
Control variables	Yes	Yes	Yes	Yes	Yes
Observations	2352	2304	2304	1818	1832

Each regression also includes a set of controls, which uses household and individual characteristics. Huber–White standard errors are reported in parentheses.
ADL, activities of daily living.
*Significant at 5% level.
**Significant at 1% level.
+Significant 10% level.

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Table III. First-stage results (dependent variable, number of children)

Instrument list	(1)	(2)	(3)
Girl*rural*73	0.797**		
	(0.053)		
Girl*rural*76			0.781**
			(0.056)
Girl*rural*79		0.793**	
		(0.060)	
Observations	2194	2194	2194
R^2	0.398	0.382	0.389
First-stage F-statistics	96.03	87.43	84.59

Each regression also includes a set of controls, which uses household and individual characteristics. Huber–White standard errors are reported in parentheses.
*Significant at 5% level.
**Significant at 1% level.
+Significant 10% level.

Table IV. Estimates of the effects of children on parental health

Health status	Girl*rural*1973	Girl*rural*1976	Girl* rural*1979
Self-reported health	0.231** (0.066)	0.132+ (0.070)	0.186* (0.075)
ADL	0.007 (0.010)	0.003 (0.011)	0.022+ (0.012)
Mental health	0.008 (0.015)	0.017 (0.016)	0.024 (0.017)
BMI	−0.095(0.088)	−0.041 (0.094)	−0.099 (0.099)
Blood pressure	−0.021 (0.092)	0.057 (0.099)	0.136 (0.107)
Observations	2194	2194	2194

Each cell corresponds to a separate regression of the health status variable on the fertility (which is instrumented), controlling for household and individual characteristics. Huber–White standard errors are reported in parentheses.
ADL, activities of daily living.
*Significant at 5% level.
**Significant at 1% level.
+Significant 10% level.

of children has no significant effect on ADL, mental health, BMI or blood pressure. These results are robust across a number of possible instruments for the number of children. The only exception is one instance when health is measured using ADL. When the number of children is instrumented using *Girl*Rural*1979*, respondents with more children have more difficulties in performing activities associated with daily living at the 10% level.

Table V shows the results where we consider the effects of the number of children on mothers' health and fathers' health separately. The rationale for distinguishing between mothers' and fathers' health is that it is possible

Table V. Effects of children on mothers' and fathers' health

	Girl*rural*1973		Girl*rural*1979	
Health status	Mothers	Fathers	Mothers	Fathers
Self-reported health	0.214* (0.094)	0.253** (0.095)	0.163 (0.111)	0.215* (0.103)
ADL	0.000 (0.013)	0.014 (0.015)	0.025 (0.016)	0.019 (0.017)
Mental health	−0.007 (0.024)	0.019 (0.019)	0.007 (0.028)	0.041+ (0.021)
BMI	0.070 (0.122)	0.110 (0.130)	−0.097 (0.141)	0.283+ (0.146)
Blood pressure	−0.035 (0.127)	0.009 (0.132)	0.184 (0.151)	0.097 (0.152)

Each cell corresponds to a separate regression of the health status variable on the fertility (which is instrumented), controlling for household and individual characteristics. Huber–White standard errors are reported in parentheses.
ADL, activities of daily living.
*Significant at 5% level.
**Significant at 1% level.
+Significant 10% level.

that mothers will experience more direct health problems later in life as a result of having children. For example, obesity later in life may be more prevalent among mothers than among fathers (Weng *et al*., 2004), as may cardiovascular disease (Zhang *et al*., 2009). There is also evidence that parent–child strains have a greater effect on mothers than on fathers, with possible adverse effects on mothers' mental health later in life (Ward, 2008).

The results in Table V are similar to the more aggregated picture in Table IV when the number of children is instrumented using *Rural*Girl*1973*. However, when the number of children is instrumented using *Rural*Girl*1979*, fathers with more children experience more adverse health effects than mothers. To be specific, fathers with more children have lower levels of self-reported health and mental health and are more likely to be obese, although the coefficient is only weakly significant in the latter two instances.

Overall, the results are fairly similar across instruments. The only instrument that gives a different result is *Rural*Girl*79* for ADL (in Table IV), self-reported health, BMI and mental health (in Table V). However, for ADL (in Table IV), BMI and mental health (in Table V), the coefficient when using *Rural*Girl*79* the instrument is only significant at the 10% level. Nonetheless, this raises the following question: why does this one instrument suggest more adverse health outcomes from having more children than the other instruments?

As our instruments correspond to different periods for the introduction of different population policies, it is not surprising that the results show some differences across IVs. According to Imbens and Angrist (1994), and Angrist *et al*. (1996), IV estimates measure the effects on outcomes for those sub-populations who are affected by the instrument. These authors term the resulting estimate the local average treatment effect (LATE) and identify this subgroup of units as 'compliers'. In our context, if we had a binary treatment indicator (such as whether someone has a second child or not) instead of the number of children, then the interpretation of LATE would be straightforward. In that case, LATE would be the average effect of having a second child on the health of parents who would decide to have a second child if the one-child policy was relaxed. When the 'number of children' is the variable of interest, LATE can be defined in a similar way: it is the weighted average marginal effect of having an additional child on parents' health should the one-child policy be relaxed.[3] The IV estimator exploiting more than one instrument is the average of the various single-instrument LATE estimators that we would obtain using each instrument separately. The weights are proportional to the effect of each instrument on the treatment variable: the larger the impact of the instrument on the regressor, the more weight it receives in the IV estimation (Angrist and Imbens, 1995).

It is very likely that non-compliance exists: not all parents who were subject to the one-child policy would have had more than one child otherwise. It is also likely that the other population control polices that were put in place in 1973 and 1976 had disparate effects on different sub-populations. Like Imbens and Angrist (1994) and Angrist *et al*. (1996), we can say that separate subgroups of units as 'compliers' exist in our case, depending on the instrument we use (the period and policy we consider). Hence, it is not surprising that *Rural*Girl*1979* gives results that are different to those from the *Rural*Girl*1976* and *Rural*Girl*1973* instruments, because each of the instruments measures the effects of having one additional child on a different subgroup of the population. It is very likely that the one-child policy, introduced in 1979, affected different groups of parents and had a larger effect, in terms of reducing the number of children, than previous policies. Thus, the impact of the one-child policy, introduced in 1979, on the effect of the number of children on parental health is likely to be larger than that of the policy changes from either 1973 or 1976.

5.4. The presence of adult children and the effect of financial transfers

The instrumented fertility effect of parental health in old age captures the net of the positive and negative effects. In this section, we seek to provide more evidence on two of the mechanisms that may potentially drive this overall net effect.

[3]This interpretation of local average treatment effect also applies in the case of non-binary instrumental variables (IVs) and non-binary endogenous regressors (Frolich, 2007). In our case, when D is the number of children, the compliance intensity can differ among households. Hence, a change in the instrument induces a variety of different reactions in D, which cannot be disentangled. Only a weighted average of these effects can be identified. For more on this issue, see Frolich (2007).

First, we investigate whether parental health is affected by adult children (over 18 years of age) living in the parents' home or in the same locality/village. The level of support provided by children is likely to be greater for adult children who are living at home or in close geographical proximity to their ageing parents. The results are reported in Table VI. Those given in the first column do not control for endogeneity. They suggest that if adult children are living with, or in close geographical proximity to, their parents, the parents' mental health is generally better, but, at the same time, they are more likely to be obese. The latter result is one instance in which the extra care provided by children results in a poorer health outcome. The presence of adult children living at home, or in close geographical proximity, has no effect on the other health variables.

In Table VI, we also investigate whether the amount of upstream financial transfers has an effect on parental health. Previous studies have suggested that upstream financial transfers are substantial in East Asia, with about 60–70% of parents receiving money from their adult children (Lillard and Willis, 1997). The results, without controlling for endogeneity, are in column (2). We find positive relationships between upstream transfers and parental health for most health indicators. Specifically, upstream financial transfers have a positive effect on self-reported health, the ease of performing day-to-day activities and mental health, although there is an adverse effect on blood pressure. Ravillion and Dearden (1988) find that transfers on the Indonesian island of Java are generally targeted towards the elderly parents who have a pre-existing illness, whereas several other studies for East Asian countries have also found parental health to be a significant predictor of upstream financial transfers (see Logan and Bian, 2003, for a review). Hence, one explanation of the result for blood pressure is that parents who receive financial transfers might have a pre-existing illness, which manifests itself in the form of having high blood pressure, and hence, causation runs in the opposite direction.

To take potential reverse causality into account, we instrument both the financial transfer and adult child living with or near parents variables separately by an interaction of the gender of the first child and his/her marital status.[4] Specifically, our instrument is whether one has a first son who is married (=1) or not. The intuitive justification for using this as an instrument is that, in China, the main responsibility for taking care of parents lies with the first married son. In western contexts, women have been described as *kinkeepers*, who feel stronger family obligations, maintain family bonds and are most involved in assistance and care giving (Ward, 2008). Such is not the case in China. A feature of many Asian societies, including China, is that parents expect a son to get married when he is self-sufficient or has a 'good' job. As parents age, they seek support from their son, both financial and non-financial. On the other hand, when a daughter marries, she usually takes care of her husband's family. Moreover, in many Asian societies, a son feels a greater responsibility towards his parents after getting married. Because he is able to provide financially for his wife and children, he will often feel that he should also be able to provide financially for his ageing parents. One survey of older residents in Beijing found that almost two-thirds (64%) agreed with the view that 'having sons makes one's old age secure' (Chen and Silverstein, 2000). To summarize, having a married son is a good instrument for co-residence in China because parents traditionally live with their oldest son and his spouse. Having a married son is a good instrument for financial transfers because financial transfers from married sons to elderly parents are much higher than from married daughters to elderly parents in both Taiwan (Lee *et al*., 1994) and China (Yang, 1996). However, one would not expect the children's marital status to have any direct effect on the parents' health in their old age.

Weak instruments tend to bias IV estimates towards OLS estimates and may weaken standard tests for endogeneity. The existing econometric literature defines the weakness of instruments based on the strength

[4]Some recent studies have attempted to test the sensitivity of the IV estimates when allowing for selection based on unobservables or when the exogeneity of the IV(s) is violated, such as Altonji *et al*. (2005) and Conley *et al*. (2012). However, we are dealing here with three types of outcome variables, that is, binary (BMI and blood pressure), categorical (self-reported health) and censored (activities of daily living and mental health); and two types of endogenous variables, binary (living with adult child) and censored (amount of financial transfer). Unfortunately, these methods cannot be applied in our context to either check the robustness of our estimates or obtain the bounds of the coefficient of interest. Moreover, these methods also set the parameters based on ad-hoc assumptions, such as the degree of correlation between unobservables that impact on parental health and financial transfer or living with adult child variables (in the case of Altonji *et al*., 2005) or prior information about the magnitude of the direct effect of child education on parental health; hence, they do not provide conclusive evidence about the range of values of the final coefficient estimates.

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Table VI. Effects on health of financial transfers and having adult children living at home, or in close geographical proximity

Health status	Results without instrumenting		IV estimates			
			First child son*marital status IV		Lewbel (2012) IV	
Self-reported health	(1)	(2)	(3)	(4)	(5)	(6)
Living with/near adult child	−0.060 (0.049)		−0.207 (0.291)		−0.219 (0.177)	
Amount of transfer		−0.02963** (0.01083)		−0.0291** (0.011)		−0.0164⁺ (0.00967)
ADL						
Living with/near adult child	−0.000 (0.007)		0.002 (0.043)		−0.00576 (0.0168)	
Amount of transfer		−0.0035** (0.0012)		−0.0040** (0.0013)		−0.0017** (0.00041)
Mental health						
Living with/near adult child	−0.025* (0.011)		0.101 (0.068)		0.0264 (0.0349)	
Amount of transfer		−0.0049** (0.0012)		−0.0048** (0.0013)		−0.0039** (0.00091)
BMI						
Living with/near adult child	0.220** (0.069)		0.746⁺ (0.386)		0.00386 (0.0837)	
Amount of transfer		0.0013 (0.0113)		0.0006 (0.0115)		0.0053 (0.0046)
Blood pressure						
Living with/near adult child	0.103 (0.071)		0.417 (0.409)		0.0267 (0.0828)	
Amount of transfer		0.0222* (0.019)		0.0183⁺ (0.0110)		0.0027 (0.0044)
First-stage			Living with/near adult child	Financial transfer		
First child son*marital status (IV)			[0.5880]** (0.0744)	[0.2660]** (0.0697)		
First-stage *F*-statistic			41.54	15.97		

Regressions control for household and individual characteristics. Huber–White SEs are given in parentheses.
ADL, activities of daily living; IV, instrumental variable.
*Significant at 5% level.
**Significant at 1% level.
⁺Significant 10% level.

of the first-stage equation. Accordingly, we test the relevance and validity of the instruments. Specifically, we test whether the IVs are correlated with the endogenous regressor and orthogonal to the error process. We test the first condition by examining the fit of the first-stage reduced-form regression of the endogenous variables on the full set of controls and IVs. The instruments are both individually significant as is evident from the first-stage reduced-form regression coefficient estimates and the corresponding standard errors, shown at the bottom of Table VI. We use the *F*-test of the joint significance of the excluded instruments in the first-stage regression. The first-stage *F*-statistics, at the bottom of Table VI, are above 15 for financial transfers and above 40 for living with adult children.

We further check the relevance of the instruments using the 'partial R^2' measure proposed by Shea (1997), which takes inter-correlations among the instruments into account. For a single instrument, Shea's partial R^2 and the usual partial R^2 measures should be the same (Baum *et al.*, 2003). In our models, Shea's partial R^2 ranges from 0.057 to 0.064. The *p*-values for Shea's partial R^2 indicate that our instruments pass the criteria recommended by Bound *et al.* (1995) and Shea (1997). Because we only have a single instrument, we cannot test the orthogonality condition of the instrument, as in the over-identification test based on the commonly used *J*-statistic of Hansen (1982).

The second-stage results are reported using the residuals from the first stage as an additional regressor. In column (3), living with, or in close geographical proximity to, a child has no effect on parental health, apart from a statistically weak effect on BMI. The results for the amount of transfers are similar to the results without instrumenting. Upstream financial transfers have been shown to be a potential source of parent–adult child strain in cases where the parents and children are bargaining over the amount or the children feel resentful of familial obligations (Ward, 2008). If the amount of upstream financial transfers is a potential source of strain, this could be the reason for the finding that parents who receive financial transfers have higher blood pressure, although the amount of financial transfers is only weakly significant for blood pressure in column (4).

Although we have attempted to show that marital status interacted with gender is a valid instrument, doubt may still linger about the exclusion restriction, which cannot be tested directly. Thus, we check the robustness of our IV results using the methodology recently proposed by Lewbel (2012), which is an identification strategy that does not rely on external instruments but rather constructs an internal IV based on the presence of heteroskedasticity in the data. The estimation problem can be summarized as

$$Y_1 = X'\beta_1 + Y_2\,\gamma_1 + \epsilon_1 \qquad \epsilon_1 = \alpha_1 U + V_1 \tag{3}$$

$$Y_2 = X'\beta_2 + \epsilon_2 \qquad \epsilon_2 = \alpha_2 U + V_2 \tag{4}$$

Let Y_1 be the health outcome and Y_2 be either living with an adult child or amount of financial transfers. U denotes the individual's unobserved characteristics, which affect both his/her health and the living arrangements or financial transfers from children. V_1 and V_2 are idiosyncratic errors. Some of the structural parameters in the aforementioned equations are not identifiable without additional information. Generally, one obtains identification by either imposing equality constraints on the coefficients of X (i.e. OLS regression) or assuming that one or more of the elements of β_1 are equal to zero. This permits the estimation of the Y_1 equation using IVs. Alternatively, assume that Z is a vector of observed exogenous variables. Lewbel (2012) argues that, if the following moment conditions are met

$$E(X\epsilon_1) = 0, E(X\epsilon_2) = 0, \quad Cov(Z, \epsilon_1\epsilon_2) = 0,$$

and there is some heteroskedasticity in ϵ_j, one can estimate the set of equations aforementioned by using the generalized method of moments with an estimate of $[Z - E(Z)]\epsilon_2$ as the instrument. The Breusch–Pagan test for heteroskedasticity rejected the null of a constant variance in each case, which is a precondition for the Lewbel (2012) strategy. We report the Lewbel (2012) estimates in the final two columns of Table VI. The results are similar to those that were obtained when using gender interacted with marital status as an IV. The only differences are that both living with or near an adult child, which had a weakly significant effect on BMI, and the amount of transfers, which had a weakly significant effect on blood pressure when using gender interacted with marital status as the IV, become insignificant.

6. CONCLUSION

There are trade-offs in the effect of fertility on parental health in old age. It is believed that the fertility of parents in developing countries is influenced greatly by the need for support in old age. Here, the focus is clearly on the benefits of having more children for receiving emotional and financial support in old age, which would generally be expected to have a positive effect on health. At the same time, increased fertility can potentially have a negative effect on parental health that persists into old age. Our main estimates of the instrumented fertility effect on parental health in old age capture the net of the positive and negative effects. We have also examined the roles of co-habitation of adult children and financial transfers to provide evidence on two of the mechanisms that drive this net effect. The results suggest that having more children has a negative effect on the self-reported parental health but generally no effect on other measures of health. The results also show that upstream financial transfers have a positive effect on parental health.

The combination of longer lives for the elderly parents and fewer children, as a result of China's one-child policy, has raised serious concerns about who is going to care for China's ageing population in the future. Jiang (1995) estimates that longer life expectancies and fewer children will quadruple the burden of caring for the elderly parents in China in the second half of the 21st century. However, an only child may feel more altruistic towards his or her parents and be more willing to make transfers to parents in their old age, consistent with the altruistic motive (Barro, 1974; Becker, 1974). Alternatively, with fewer children, the well-known quantity–quality trade-off suggests that parents will invest more in their children. As a consequence, children may be more willing to repay their parents with transfers as their parents age, consistent with the exchange motive (Bernheim *et al.*, 1985; Cox, 1987). The increasing incomes resulting from China's rapid economic growth have increased the opportunity cost of having adult children and been a catalyst for the breakdown in traditional family structures (Logan and Bian, 2003). In these circumstances, children tend to make financial transfers to their parents in lieu of time transfers. Our findings suggest that upstream financial transfers are an effective means of addressing concerns for the well-being of China's ageing population in the face of the declining number of children.

A limitation of our results is that we rely on cross-sectional data from the CHARLS pilot study. Although we have used a range of identification strategies to show that our results are robust, doubt may exist about the exclusion restriction on our conventional IVs, which is not testable. We find similar results using method suggested by Lewbel (2012). The robustness using this alternative method, which does not rely on exclusion restriction, gives us added confidence about our IV results. The Lewbel (2012) approach, however, depends on assumptions about the nature of the heteroskedasticity in the data, which are not testable. Future research can use other Chinese datasets focusing on ageing and health or similar health and retirement data from other countries to examine whether such results hold in the context of other developing countries.

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