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# Number of Children and Living Arrangements of the Elderly in China

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ABSTRACT

This study examines the effects of the number of children on the living arrangements of married and widowed individuals using data from the 2008 wave of the Chinese Longitudinal Healthy Longevity Survey. While the literature offers mixed results owing to the endogeneity bias from the number of children, we use sex dummy variables of the first and second parity as instrumental variables to correct the bias and find that the number of children has a statistically significant and positive effect on the probability of cohabitation for both subsamples. The magnitude of this effect in the widowed subsample is about 1.64 times that in the married subsample, indicating that children play a more important role in the case of widowed individuals.

Keywords: living arrangements, one-child policy, China, marital status, elderly population.

JEL classification: D13, J1

## 1. Introduction

With medical technology innovations and the resultant extension in lifespan, China is experiencing a dramatic increase in its aging population. An increasing rate of its elderly population is reorganizing their living arrangements because of changing individual characteristics, such as the deterioration of their functional ability. Faced with an underdeveloped social security system, China's elderly population is generally dependent on their adult children to provide physical, emotional, and financial support (Korinek, Zimmer, & Gu, 2011). This situation is more adverse in China compared to other developing and developed countries due to the lack of a mature social security system and its large aging population. Moreover, majority of China's elderly population, especially in the rural areas, live below the poverty line. A factor that further impeded traditional routes of elderly support is the government's one-child policy, although it was later relaxed in 2015 to a maximum of two children per couple. While such policy measures may influence the decision of childbirth and relieve the burden of elderly support in the long-run, predicting the short-run effects is difficult. Moreover, given the dramatic speed at which the elderly population is increasing, we lack sufficient time to confirm related policy effects in China.

The effects of the number of children on one's living arrangements in old age have gained much attention from policymakers and economists alike. However, since the one-child policy was implemented in 1979, the elderly who were completely affected by it are too young to make living arrangements. Thus, we use data for individuals aged 65 year and older to predict the effects of the number of children on living arrangements of the elderly keeping other factors constant.

Several studies examine the determinants of living arrangements of the elderly (Hoerger, Picone, & Sloan, 1996; Guo, 2002; Meng & Luo, 2008; Ren & Treiman, 2014; Zhang, 2015), although most of them focus on economic factors such as pension benefits and bequest motives (Edmonds, Mammen, & Miller, 2004; Engelhardt et al., 2005; Meng & Luo, 2008; Yin, 2010) and include number of children as a control variable. While researchers find that the number of children positively affects the probability of cohabitation (Börsch-Supan, Haijvassiliou, Kotlikoff, & Morris, 1992; Hoerger et al., 1996; Ren & Treiman, 2014), some others report negative effects (Guo, 2002; Ren & Treiman, 2014; Zhang, 2015). The literature, thus, lacks reasonable explanations owing to the adverse causality between number of children and living arrangements of the elderly. Hence, in this study, we focus on the causal effects of the number of children on living arrangements of the elderly using instrumental variable design to correct the endogeneity bias. To the best of our knowledge, this study is the first to examine this causal relationship by correcting the endogenous bias.

This study estimates the effects of the number of children on the living arrangements of married and widowed individuals using an instrumental variable design to eliminate any endogenous bias arising from the number of children. Angrist and Evans (1998) accounted for parental preference for a mixed sibling-sex composition to construct instrumental variables to estimate the effects of childbearing on labor supply. Following their study, we correct the endogeneity bias of the number of children using the preference for a male child.

We use data from the 2008 wave of the Chinese Longitudinal Healthy Longevity Survey and find that the number of children has a statistically significant and positive effect on the probability of cohabitation for both subsamples. The magnitude of the effect in the widowed subsample is about 1.64 times that in the married subsample, indicating that children play a greater role in the case of widowed individuals. This result is contrary to that in the extant literature that does not control for the endogeneity bias.

The remainder of this paper is organized as follows. Section 2 describes the data and summary statistics. Section 3 presents our hypothesis. Sections 4 and 5 discuss the empirical strategy to estimate causal effects and the results. Section 6 provides concluding remarks and policy implications.

## 2. Data

### 2.1 Data

Data used in this paper are from the 2008 wave of the Chinese Longitudinal Healthy Longevity Survey (CLHLS). CLHLS data were collected by Peking University Center for Healthy Aging and Family Studies and China Mainland Information Group. The baseline survey was administered in 1998 to a randomly selected number of counties and cities from China's 22 provinces: Liaoning, Jilin, and Heilongjiang (northeast); Beijing, Tianjin, Hebei, and Shanxi (north); Shanghai, Jiangsu, Zhejiang, Anhui, Fujian, Jiangxi, and Shandong (east); Henan, Hubei, Hunan, Guangdong, and Guangxi (central/south); and Chongqing, Sichuan, and Shaanxi (west). An enumerator and a nurse or medical school student conducted the interview and a basic health examination at each interviewee's home.

The total number of interviewees in the 2008 wave was 16,540, including 341 adult children younger than 65 years of age. Since we focus on the living arrangements of the elderly, we exclude adult children from our analysis. In addition, given our objective to identify the causal effects of the number of children on living arrangements of the elderly, we use an over-identified model and exclude the elderly with one or no children.

### 2.2 Dependent Variable

The dependent variables measure three types of living arrangements for the elderly: living with family members, independently, or in institutions for the elderly. Family members include spouse, children, relatives, and parents, and a majority of the total sample

(83.25%) reportedly lived with their family. Given the focus of our study, we exclude respondents who lived with relatives and parents because of systematic difference from cohabitating with children. About 14.89% of the sample stated they lived independently, which means to live alone without a spouse or any other family member. Finally, only 1.86% of the sample lived in an institution. We excluded such respondents because most of them receive financial support from China's local governments.

Some researchers classify living with a spouse as "living independently" (Hoerger et al., 1996; Meng & Luo, 2008; Zhang, 2015). Studies have shown a significant rise in cohabitation following widowhood (Korinek et al., 2011), indicating that marital status is an important determinant of living arrangements. In addition, sex preferences for children differ by generation and we cannot use cross-sectional data to control for such preferences. Table 1 presents the summary statistics of the variables. As shown in the table, the average age of the married and widowed subsamples ranges from 78.197 to 91.533. It is possible to control for cohort effects from age using different subsamples; thus, we create two subsamples on the basis of marital status.

The married subsample includes respondents with a living spouse at the time of the survey. We define the dependent variable, cohabitation, in the married subsample as follows. Respondents living with a spouse and children are assigned the value of one and zero otherwise. We apply a similar definition of cohabitation for the widowed subsample, which includes respondents who were widowed during the survey. Respondents living with their children were assigned a value of one and zero otherwise. We selected 4,102 and 9,491 valid respondents for the married and widowed subsamples. Unmarried respondents were excluded from our analysis. Although we use cohabitation as a dependent variable for married and widowed individuals, its contents differ between both subsamples.

Table 1 shows a wide variation between the two subsamples in terms of cohabitation (married: 30.6% and widowed: 78.2%), suggesting that marital status is an important determinant in the living arrangements of the elderly (Hoerger et al., 1996; Zhang, 2015). The elderly tend to live with their children following the death of their spouse. This is reasonable since the elderly need physical support owing to deterioration in functional abilities or emotional support in the case of a deceased spouse (Korinek et al., 2011).

### 2.3 Independent Variables

In China, public pension and social welfare systems remain underdeveloped and the elderly primarily depend on their children for physical, emotional, and financial support (Korinek et al., 2011). Before the 1970s, China mainly depended on agriculture. The only way to improve output and increase their wealth was to extend their labor force by fostering several children. Consequently, the average number of children (Children) in our subsamples is more than four (married: 4.510 and widowed: 4.925). The high average number of children should be explained with some caution. To employ an over-identified model, we exclude individuals with one child or no children, which seems to increase the sample average. However, those who were subject to the one-child policy had one child, unless they were exempted (Zhu, 2003). Thus, we see an obvious decline in the average number of children before and after the one-child policy.

We control for age (Age) and the quadratic form of age (Age_2) to capture the nonlinear trend of cohabitation. According to previous studies, the sign of age can be modified before and after a certain threshold. We expect the sign of age to be negative and that of its quadratic form to be positive. This is because the elderly tend to need less physical or emotional support before the threshold and live independently; however, as they grow older, these needs are likely to increase, especially after the threshold. As mentioned, the age widely varies between the two subsamples (married: 78.197 and widowed: 91.533). The rate of elderly mortality rapidly grows with an increase in age and thus, the average age in the widowed subsample is higher than that in the married subsample.

We believe activities of daily living (ADL) are the most appropriate instrument to measure elderly health status given the tasks accounted for (bathing, dressing, toileting, indoor transferring, feeding, and continence). We use a dichotomous variable to measure each of the six items (respondents are assigned a value of one when they need assistance and zero otherwise) and calculate the total score, ranging from zero to six, where zero means no limitation on ADL. The two subsamples considerably vary in ADL (married: 0.229 and widowed: 0.839), indicating that the respondents in the married subsample are healthier than those in the widowed subsample. This may also imply that the probability of cohabitation in the married subsample is lower than that in the widowed subsample owing to the greater need for physical support in the latter. However, the causal explanation of ADL may be subject to the endogeneity bias in health conditions; thus, one should be cautious when explaining the causal effects of ADL when not accounting for endogeneity. In this study, however, we include ADL because our primary aim is to minimize omitted variable bias rather than interpret the causal effects of ADL estimates (Van Houtven & Norton, 2004).

We use annual household income in the previous year (Income), ownership of real estate property (Ownership), and main financial support (Financialrt, Finacialch, Financiallg, and Financialwb) to control for the effects of economic factors. We find no significant difference in the average annual household income of the previous year[1] between the two subsamples (married: 22.030 and widowed: 22.573; unit: 1,000 yuan). In the case of a higher annual household income, the elderly may be economically independent and live away from their children and possibly with their spouse. Since we

[1] According to CLHLS' coding book, if the annual household income in the previous year was higher than 100,000 yuan, the final data recorded it as 100,000 yuan. We made the following rectification to correct such data bias: if the respondent reported an annual income of 100,000 yuan, we recorded it as 125,000 yuan.

cannot clearly determine contribution by an elderly member to annual household income, we explain income with some caution. The rate of real estate owners in the married subsample is about three times that in the widowed subsample but only 14.4% on average. We expect that this economic factor is a good proxy for bequest motivation. If the elderly own real estate, it is likely that their adult children will cohabitate with them to inherit the property by providing informal care (Yin, 2010). The main source of financial support includes retirement pension, adult children, local government aid, self-employment, and other support measures. We use binary variables to define main sources of financial support. If the main financial source is retirement pension, Financialrt takes the value of one and zero otherwise. We treat Financialch (adult children), Financiallg (local government), Financialwb (self-employment), and other sources as reference groups. The most important source of financial support in our subsamples is adult children (married: 42.3% and widowed: 78.6%) (Korinek et al., 2011). In addition to annual household income, we use retirement pensions (Financialrt) to measure the degree of economic independence (married: 29.4% and widowed: 11.3%).

Sex is measured dichotomously: the variable takes the value of one in the case of a male respondent and zero for a female respondent. The rate of male respondents in the married subsample (66.9%) is more than twice that in the widowed subsample. This result can be attributed to the shorter life span of males. The rural variable is also dichotomously coded on the basis of the respondents' residence in 2008: the respondent takes the value of one if he/she is from a rural area or town and zero otherwise. However, there is no significant difference between the two subsamples in terms of residence (married: 57.9% vs. widowed: 62.7%). We also control for ethnicity (ethnic = 1, non-ethnic = 0). Education is measured by a continuous variable. Most of the elderly in our sample did not have the opportunity to enroll in school and therefore, educational attainment levels are rather low. Table 1 shows that educational attainment for the married subsample is three times that for the widowed subsample (3.434 years). Finally, we control for variation in provinces (reference group: Shaanxi), which are not shown in Table 1.

Table 1 Summary Statistics

| | Married | | Widowed | |
|---|---|---|---|---|
| Variables | Mean | Standard deviation | Mean | Standard deviation |
| Cohabitation | 0.306 | 0.461 | 0.782 | 0.413 |
| Children | 4.510 | 1.764 | 4.925 | 2.037 |
| Age | 78.197 | 9.134 | 91.533 | 9.656 |
| Age_2 | 6198.255 | 1472.594 | 8471.548 | 1717.555 |
| Sex | 0.669 | 0.471 | 0.300 | 0.458 |
| Rural | 0.579 | 0.494 | 0.627 | 0.484 |
| Ownership | 0.140 | 0.347 | 0.041 | 0.199 |

| | | | | |
|---|---|---|---|---|
| Ethnicity | 0.039 | 0.194 | 0.055 | 0.228 |
| Income | 22.030 | 26.649 | 22.573 | 26.507 |
| ADL | 0.229 | 0.931 | 0.839 | 1.651 |
| Education | 3.473 | 4.060 | 1.261 | 2.706 |
| Financialrt | 0.294 | 0.456 | 0.113 | 0.316 |
| Financialch | 0.423 | 0.494 | 0.786 | 0.410 |
| Financiallg | 0.035 | 0.183 | 0.052 | 0.222 |
| Financialwb | 0.176 | 0.381 | 0.033 | 0.177 |
| Observations | 4,102 | | 9,491 | |

Note: Income is presented in units of 1,000 yuan.

## 3. Hypothesis

### 3.1 Hypothesis

Economic theory of fertility indirectly determines fertility as a result of parental choice between current and future consumption; thus, children can be thought of as an investment good (Cochrane, 1975). This is particularly the case in developing countries, where public pension and social welfare systems are largely rudimentary. The elderly have to depend on the support of their adult children in their later life. However, previous studies offer mixed results on the effects of number of children on cohabitation with the elderly. Thus, this study attempts to provide empirical evidence on the topic by correcting the endogeneity bias. We consider cohabitation as the most efficient way of gaining physical and emotional support as it minimizes transaction costs such as travel time (Hoerger et al., 1996). Accordingly, we hypothesize that, ceteris paribus, an additional child increases the probability of cohabitation.

## 4. Empirical strategy

### 4.1 Identification strategy

This study aims to examine the causal effects of the number of children on living arrangements of the elderly. However, using an OLS or probit model may give us biased estimation results owing to the endogeneity from the number of children. For example, the elderly may try to obtain sufficient support and accordingly, decide the optimal number of children, indicating the possibility of a reverse causal relationship. We draw on Angrist and Evans (1998), who use parents' sex preferences and mixed sex preferences regarding children as instrumental variables to correct the endogeneity bias and estimate the impact of number of children on parents' labor supply. As China mainly depended on agriculture prior to the 1970s, there were strong preferences for a son who could improve output and join the labor force. We assume a direct influence of preference for a male child on the number of children.

To correct the endogeneity bias, we apply an instrumental variable approach using male dummies for the first and second parities. We use a dichotomous variable to measure each instrumental variable. Respondents with a son are assigned the value of one and zero otherwise. We assume that the bearing of a male child decreases the total number of children because parents' demand for a male child is satisfied. Thus, the signs of both instrumental variables are assumed to be negative.

The following regression models are used to link the number of children and our instrumental variables:

$$\text{Children}_i = \alpha_1 + \beta_1 \text{firstparity}_i + \beta_2 \text{secondparity}_i + X\lambda + \eta_i, \qquad (1)$$

where $\text{Children}_i$ denotes number of children, subscript i is respondent number, $\text{firstparity}_i$ and $\text{secondparity}_i$ are dummy variables indicating male of first and second parities, $\lambda$ is a coefficient vector, and $\eta_i$ is an error term. $X$ is a vector for control variables including age, square of age, sex dummy, ethnicity dummy, rural area dummy, ADL, educational attainment, ownership of real estate property dummy, annual household income in the previous year, main financial support dummies, and prefectures dummies. We expect the signs of $\beta_1$ and $\beta_2$ to be negative, indicating that preferences for a son decrease the optimal number of children.

In the first stage, we use equation (1) to predict the number of children. In the second stage, we use equation (2) to estimate the effects of the number of children on living arrangements for the elderly:

$$Y_i = \alpha_2 + \beta_3 \text{PChildren}_i + Z\delta + \varepsilon_i, \qquad (2)$$

where $Y_i$ denotes living arrangements of the elderly defined using dichotomous variables (cohabitation: Yes = 1, No = 0), subscript i refers to respondent number, $\text{PChildren}_i$ is the predicted number of children using equation (1), and $\beta_3$ is the causal effect of our interest. We expect $\beta_3$ to be positive, indicating that an additional child increases the probability of cohabitation. $\delta$ is a coefficient vector and $Z$ is a vector for control variables including age, square of age, sex dummy, ethnicity dummy, rural area dummy, ADL, educational attainment, ownership of real estate property dummy, annual household income in the previous year, main financial support dummies, and prefectures dummies. $\varepsilon_i$ is an error term.

## 5. Estimation results

### 5.1 Estimates of children equation

Table 2 presents the first stage results of the children equations for the two subsamples. Column 1 shows negative first (male) and second (male) parities for the married subsample, indicating a decrease in the total number of children without controlling for covariates, which is consistent with our expectation (first parity: −0.226 and second parity: −0.178). We see that the magnitude of coefficients in column 2 reduces compared to that of coefficients in column 1. However, the results show that the coefficients of the two variables are negative and statistically significant at the 1% level (first parity: −0.194 and second parity: −0.167). The results for the widowed subsample are presented in columns 3 and 4. The signs of the coefficients for the first and second parity dummies in column 3 are both statistically significant and negative without controlling for covariates. However, the coefficient of the second parity (male) is statistically non-significant at the 10% level when controlling for covariates[2]. The magnitude of all coefficients in the widowed subsample is less than that of coefficients in the married subsample, indicating that the effects of preference for a son on childbearing differ between the two subsamples, possibly because of a cohort effect. Overall, our results are consistent with the hypothesis that a male child (first and second parities) decreases the total number of children.

Table 2 Estimates of Children Equations

| | Married | | Widowed | |
|---|---|---|---|---|
| Independent variable | (1) | (2) | (3) | (4) |
| First parity | −0.226*** | −0.194*** | −0.099** | −0.101** |
| | (0.055) | (0.049) | (0.042) | (0.042) |
| Second parity | −0.178*** | −0.167*** | −0.079* | −0.067# |
| | (0.050) | (0.049) | (0.042) | (0.041) |
| Other covariates | No | Yes | No | Yes |
| Adj $R^2$ | 0.0061 | 0.2149 | 0.0007 | 0.0427 |
| Obs. | 4,062 | 4,062 | 9,360 | 9,360 |

Notes:

Dependent variable: number of children

*, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively. # represents statistical significance at the 10.3% level.

The figures in parentheses are standard errors.

## 5.2 Married subsample

Columns 1–3 in Table 3 present the estimation results for the married subsample. Column 1 shows the probit model estimates for the probability of cohabitation without correcting the endogeneity bias of the number of children. The covariates include age, square of age, sex dummy, ethnicity dummy, rural area dummy, ADL, educational attainment,

[2] We find that the coefficient of the second parity (male) is statistically significant at the 10.3% level.

ownership of real estate property dummy, annual household income of the previous year, main financial support dummies, and prefectures dummies. It also shows that the coefficient for number of children is statistically significant and negative at the 1% level. As for the magnitude, an additional child decreases the probability of cohabitation by 1.4 percentage points, which is in contrast to our expectation. This is possibly because a higher number of children mean greater potential for daily care or financial support, thus diminishing the likelihood of cohabitation (Zhang, 2015). However, this estimation result may be contaminated by the endogeneity bias. To correct this, we use an instrumental variables approach and find that the coefficients for number of children becomes positive and statistically significant at the 5% level (columns 2 and 3 of Table 3), which is in line with our expectations. We also find that, after controlling for effects from other factors, an additional child increases the probability of cohabitation by 66.6 percentage points and the marginal probability is about 10 percentage points higher than that without covariate estimates.

In terms of age (column 3 of Table 3), we find that an additional year significantly decreases the probability of cohabitation by 62.3 percentage points. When using the quadratic form of age to capture increasing effects, the quadratic must eventually turn around and the sign of the quadratic is expected to be positive. Thus, when age passes the turning point, the probability of cohabitation will increase with age. We can obtain a turning point using the coefficient for age that is more than twice the absolute coefficient value for the quadratic form of age (Wooldridge, 2012). The turning point in this study is 78. In addition, we find that males and those belonging to an ethnic minority cohabitate with a higher probability than their counterparts. The likelihood of cohabitation for the elderly living in rural areas significantly decreases by 28.1 percentage points compared to those residing in the city. Although additional ADL significantly increases the probability of cohabitation by 9.7 percentage points, the potential endogeneity of limitations in ADL may bias the causal explanation for the estimated results. Thus, one must exercise some caution with explaining ADL. The purpose of including ADL is to reduce variation in the error term and improve the preciseness of estimated coefficients. The coefficient for educational attainment shows a positively correlation with the likelihood of cohabitation but is statistically non-significant. The coefficient for retirement pension as a main financial resource is statistically positive, indicating that adult children consider retirement pension to be an attractive factor when deciding to cohabitate with their elderly parents. This result is contrary to our expectation that financial independence decreases the likelihood of cohabitation for the elderly. However, since cohabitation results from negotiation between elderly parents and their adult children, the latter may not live economically independent from their parents and have to cohabitate with them, which is consistent with the result for annual household income in the previous year. In the case of affluent parents or households, adult children prefer to live with their elderly parents. We do not find ownership of real estate property to significantly increase the probability of cohabitation, indicating that bequest motivation is too small to detect or real estate ownership may not be a good proxy for bequest motivation.

In general, two indexes are applied when validating instrumental variables. A good instrumental variable must be satisfied by two conditions. First, the instrumental variables must be highly correlated with the endogenous variable. We use first stage F-statistics to evaluate the correlation between the instrumental and endogenous variables. We derive F-statistics (columns 2 and 3) of 13.42 and 31.89, which is greater than 10, indicating that our instrumental variables are highly correlated with the endogenous variable (Stock & Watson, 2011). The second condition is that the instrumental variables are not correlated with the error term. Since we construct an over-identified model using two instrumental variables, we use Amemiya–Lee–Newey's minimum chi-square statistic to test the second condition. The null hypothesis is that instrumental variables are not correlated with the error term. The result does not reject the null hypothesis at the 5% significance level and thus, our instrumental variables are valid.

5.3 Widowed subsample

We adopt the same estimation method as that for the married subsample. We use a probit model to estimate the effects of number of children on the probability of cohabitation. The results are reported in column 4 of Table 3 and include the same covariates as those in the married subsample. The coefficient that is of interest to this study, number of children, is statistically significant at the 1% level. An additional child decreases the probability of cohabitation by 0.8 percentage point. However, it is doubtful whether this result can be interpreted as the real causal effect of number of children on the probability of cohabitation owing to the endogeneity bias.

To eliminate the bias, here as well, we use preference for a son as an instrumental variable to estimate the coefficient for number of children. The results are presented in columns 5 and 6 of Table 3. After controlling for the covariates, we find that the magnitude of the coefficients in column 6 increases to more than the magnitude of those in column 5 and the coefficients for number of children are positive and statistically significant at the 5% level. The results in column 6 show that an additional child increases the probability of cohabitation by 109.3 percentage points. Although the first stage F-statistics in column 5 are less than 10, indicating the instrumental variables may be weak instruments, after controlling for the covariates, we find that our instrumental variables are highly correlated with the endogenous variable given that the F-statistics are greater than 10.

As for age effects, we find that an additional year of age decreases the probability of cohabitation by 29.7 percentage points. We find no significant result for sex or ethnic minority. The effect of the rural dummy decreases the probability of cohabitation by 23.4 percentage points, which is a finding similar to that for the married subsample. Since ADL is highly correlated with age, the effect of ADL significantly increases the probability of cohabitation by 17.7 percentage points. Here as well, it is important to explain ADL with some caution.

Next, we find that retirement pension and annual household income statistically increase the probability of cohabitation, indicating that retirement pension and annual household income are attractive factors for adult children to cohabitate with their elderly parents. Widowed elderly may need more physical and emotional support from their children and thus, children as the main financial resource increases the probability of cohabitation, which is in contrast with our findings for the married subsample. An interesting result is that ownership of real estate property significantly decreases the probability of cohabitation by 37.9 percentage points. A possible explanation is that the probability of inheriting property increases as parents grow older and all children may want to cohabitate owing to strong bequest motives. In this case, elderly parents may find it difficult to choose which child to live with and this may influence the cohabitation behavior of the elderly.

Although the magnitudes of the estimation results for the two subsamples differ, the sign of the coefficients is consistent with our hypothesis that the likelihood of cohabitation increases with the number of children.

Table 3 Probit and IV probit estimates for living arrangement models

| | Married | | | Widowed | | |
|---|---|---|---|---|---|---|
| | (1) | (2) | (3) | (4) | (5) | (6) |
| Estimation method | Probit | IV probit | IV probit | Probit | IV probit | IV probit |
| Children | −0.014*** | 0.559*** | 0.666*** | −0.008*** | 1.031** | 1.093** |
| | (0.005) | (0.186) | (0.220) | (0.002) | (0.418) | (0.471) |
| Age | −0.081*** | | −0.623*** | −0.023*** | | −0.297*** |
| | (0.013) | | (0.130) | (0.007) | | (0.098) |
| Age_2 | 0.000*** | | 0.004*** | 0.000*** | | 0.002*** |
| | (0.000) | | (0.001) | (0.000) | | (0.001) |
| Sex | 0.018 | | 0.261*** | −0.056*** | | −0.066 |
| | (0.017) | | (0.092) | (0.0101) | | (0.090) |
| Ethnicity | 0.112*** | | 0.337** | 0.054*** | | 0.050 |
| | (0.044) | | (0.152) | (0.016) | | (0.159) |
| Rural | −0.034* | | −0.281*** | −0.019** | | −0.234** |
| | (0.018) | | (0.090) | (0.009) | | (0.092) |
| ADL | 0.035*** | | 0.097*** | 0.048*** | | 0.177*** |
| | (0.008) | | (0.030) | (0.004) | | (0.022) |
| Education | −0.001 | | 0.015 | −0.001 | | 0.008 |
| | (0.002) | | (0.010) | (0.002) | | (0.013) |
| Financialret | 0.058 | | 0.364** | 0.072*** | | 0.678** |
| | (0.035) | | (0.143) | (0.024) | | (0.281) |
| Financialch | 0.118* | | 0.192 | 0.115*** | | 0.508** |

| | | | | | | |
|---|---|---|---|---|---|---|
| | (0.032) | | (0.127) | (0.035) | | (0.221) |
| Financiallg | 0.024 | | 0.054 | −0.026 | | 0.390 |
| | (0.053) | | (0.191) | (0.036) | | (0.317) |
| Financialwb | 0.030 | | 0.074 | −0.039 | | 0.225 |
| | (0.036) | | (0.133) | (0.039) | | (0.305) |
| Ownership | −0.062*** | | −0.032 | −0.125*** | | −0.379** |
| | (0.023) | | (0.106) | (0.027) | | (0.150) |
| Income | 0.003*** | | 0.012*** | 0.002*** | | 0.007*** |
| | (0.000) | | (0.001) | (0.000) | | (0.001) |
| Prefecture | Yes | No | Yes | Yes | No | Yes |
| Obs. | 4,102 | 4,062 | 4,062 | 9,491 | 9,360 | 9,360 |
| First stage F-statistics | - | 13.42 | 31.89 | - | 4.43 | 12.6 |
| Chi-sq. statistic | - | 0.408 | 0.006 | - | 0.008 | 0.000 |
| P-value | - | 0.5229 | 0.9358 | - | 0.9301 | 0.9951 |

Note:

Dependent variable: cohabitation with children (Yes = 1, No = 0).

*, **, and *** indicate statistical significance at the 10%, 5%, and 1% levels, respectively.

"Chi-sq. statistic" represents the Amemiya–Lee–Newey minimum Chi-square statistics.

The figures in parentheses are standard errors.

## 6. Discussion and Conclusions

Studies have examined the effects of economic resources and demographic changes on living arrangements of the elderly using number of children as a control variable, but have largely provided mixed results owing to the endogeneity bias. This study, however, focuses on the causal effects of the number of children on the living arrangements of the elderly in China using an instrumental variable design to correct the endogeneity bias. In addition, we use preference for a male child as an instrumental variable, which is reasonable given that such a preference prevails in China. Our study reveals that the number of children positively affects the probability of cohabitation in China.

Cohabitation seems to be the most efficient way to support the elderly in China. However, to control the population scale, in 1979, the Chinese government imposed the one-child policy. The policy aimed to reduce the pressure of increasing population on natural resources by limiting the number of children to one per set of parents, with some exceptions. As the exogenous shock of number of children significantly affects the probability of cohabitation, parents who were influenced by the one-child policy may not obtain sufficient support from their children in their old age. In addition, weak public pension and social welfare systems may widen the gap between the demand and supply for physical, emotional, and financial support and the elderly were likely to foster many children to ensure sufficient support.

Although the Chinese government abolished the one-child policy, allowing all couples to have two children, it has not arrested its efforts to control the population scale. According to our estimation results, an additional child increases the probability of cohabitation; however, the Chinese people still cannot choose to bear their desired number of children. Since the respondents in our data are mainly 65 years or older, their childbearing behavior was not or partially affected by the one-child policy and their average number of children is greater than four. However, parents completely affected by the one-child policy may have only about one-fourth of the average number of children compared to that in our sample. In the future, we do not expect the influencing factors considered in this study (Table 1) to drop as rapidly as the number of children.

The number of children can significantly affect the living arrangements of the elderly. With the public pension and social welfare systems developing, the function of children in supporting the elderly may gradually weaken. However, the development of such systems is highly dependent on the younger generation. For example, the present “pay as you go” pension system is entirely supported by the younger generation, in the absence of which the system could collapse. Therefore, abolishing family planning policies and improving the fertility rate play a crucial role in China’s response to the increasing elderly population.

Despite its contribution, this study is subject to certain limitations. The data used in this study are from the 2008 wave of the Chinese Longitudinal Healthy Longevity Survey. Although we try to control for the cohort effect of age and use two subsamples to estimate the effects of the number of children, we cannot completely detect the cohort effect on living arrangements, which may bias the estimation results. Thus, future studies should consider applying panel data to control for the cohort effect of age and other dynamic factors.

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References


Angrist, J. D., & Evans, W. N. (1998). Children and their parents' labor supply: Evidence from exogenous variation in family size, *American Economic Review*, 88, 450-477.

Börsch-Supan, A., Haijvassiliou, V., Kotlikoff, L. J., & Morris, J. N. (1992). Health, children, and elderly living arrangements: a multiperiod-multinomial probit model with unobserved heterogeneity and autocorrelated errors. *Topics in the Economics of Aging*. http://doi.org/10.3386/w3343

Cochrane, S. H. (1975). Children as by-products, investment goods and consumer goods: A review of some micro-economic models of fertility. *Population Studies*, 29, 373-390. http://doi.org/10.2307/2173934

Edmonds, E. V, Mammen, K., & Miller, D. L. (2004). Rearranging the family? Income support and elderly living arrangements in a low-income country. *Journal of Human Resources*, 40, 186–207. http://doi.org/10.3368/jhr.XL.1.186

Engelhardt, G. V, Gruber, J., & Perry, C. D. (2005). Board of regents of the university of Wisconsin system social security and elderly living arrangements: Evidence from the social security notch. *Journal of Human Resources*, 40, 354-372.

Guo, Z. (2002). Zhongguo gaoling laoren de juzhu fangshi jiqi yingxiang yinsu (The living arrangements of the elderly and its influencing factors), *Population Research*, 14-18.

Hoerger, T. J., Picone, G. A., & Sloan, F. A. (1996). Public subsidies, private provision of care and living arrangements of the elderly, *The Review of Economics and Statistics*, 78, 428-440.

Korinek, K., Zimmer, Z., & Gu, D. (2011). Transitions in marital status and functional health and patterns of intergenerational coresidence among China's elderly population. *Journals of Gerontology - Series B Psychological Sciences and Social Sciences*, 66 B, 260-270. http://doi.org/10.1093/geronb/gbq107

Meng, X., & Luo, C. (2008). What determines living arrangements of the elderly in urban China. *Inequality and Public Policy in China*, 267-286. http://doi.org/10.1017/CBO9780511510922.011

National Health and Family Planning Commissions of the People's Republic of China (2016), "Law of the People's Republic of China on Population and Family Planning," http://www.nhfpc.gov.cn/zhuzhan/wsjsfl/201606/1c0ac828bce44560bd2389cba203663b. Accessed 16.06.20

Ren, Q., & Treiman, D. J. (2014). Living arrangements of the elderly in China and consequences for their emotional well-being. *Chinese Sociological Review*, 47, 255-286. http://doi.org/10.1080/21620555.2015.1032162

Stock, J. M., & Watson, M.W. (2011). *Introduction to econometrics.* (3rd ed.). Addison-Wesley, Boston, (Chapter 12).

Van Houtven, C. H., & Norton, E. C. (2004). Informal care and health care use of older adults.

*Journal of Health Economics*, 23, 1159-1180. http://doi.org/10.1016/j.jhealeco.2004.04.008

Wooldrige, J. M. (2012). *Introductory econometrics a modern approach.* (5th ed.). South-Western: Cengage Learning, (Chapter 6).

Yin, T. (2010). Parent-child co-residence and bequest motives in China. *China Economic Review*, 21, 521-531. http://doi.org/10.1016/j.chieco.2010.05.003

Zhang, L. (2015). Determinants of living arrangements among the Chinese elderly : New evidence from the CLHLS 2011 wave, *Chinese Studies*, 4, 32–43.

Zhu, W. X. (2003). The one child family policy. *Archives of Disease in Childhood*, 88, 463-464.