# Risk Behaviors of Only-Child Parents: Evidence from the One-Child Policy in China*

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**Abstract**

This paper examines the impact of having only one child on parental risk behaviors in a society where children serve as the bedrock for elderly support. Using variation in the One-Child Policy in China, we find that elderly parents with one child consistently show more risk avoidance behaviors in both health and finance domains. Moreover, the effect is intensified where the children-for-support culture is stronger and buffered where institutional elderly support is better available. Lastly, only-child parents also show stronger risk aversion in preferences. Together, the study underscores the role of family structure in shaping individual risk behaviors and preferences.

JEL Codes: D10, D15, D81, J13, I12, G11

Keywords: risk behavior, risk preference, only child, fertility policy, elderly support

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# 1 Introduction

Risk behaviors are essential for understanding individual economic activities and well-being. A rich set of literature has documented the determinants of individual risk behaviors and preferences, including gender, age, and education (e.g., Borghans et al., 2008; Dohmen et al., 2011; Outreville, 2015; Schildberg-Hörisch, 2018), cognitive skills and psychological traits (Dohmen et al., 2018; Mata et al., 2018), macroeconomic fluctuations (Malmendier and Nagel, 2011), health shocks (Decker and Schmitz, 2016), exposure to violence (e.g., Voors et al., 2012; Callen et al., 2014), near-miss accidents among automobile drivers (Shum and Xin, 2022), and natural disasters (e.g., Cameron and Shah, 2015; Hanaoka et al., 2018). However, considering the fundamental role of families in risk sharing, understanding the effect of family structure on risk behaviors has remained relatively scarce.[1]

This paper examines a new dimension of family configuration on individual risk behaviors, focusing on the impact of having only one child on parental risk behaviors in China. The China context has two main features that are relevant to answer the question. On the one hand, there is a long-standing culture regarding the role of children in supporting elderly parents (e.g., Yuan, 2004; Banerjee et al., 2014).[2] On the other hand, in recent decades, China has experienced restrictive fertility policies, of which the One-Child Policy has played an important role. As a result of the policy implementation and socioeconomic transition, the one-child family – parents with only one child during their lifetime – has been an increasingly important type of family structure in China (Wang, 2009; Zhang, 2017). Figure 1 illustrates the trends in China's Total Fertility Rate (TFR) and the share of only children in the population.[3] As shown by Figure 1, the proportion of only children has increased by birth cohort in recent decades, from 2% in 1965 to approximately 10% in 1990.

With changes in family structure, the tension between the deterioration of family insur-

[1] Existing studies have focused primarily on the role of parenthood on risk attitudes and behaviors (DeLeire and Levy, 2004; Grazier and Sloane, 2008; Cameron et al., 2010; Love, 2010; Kerr et al., 2011; Görlitz and Tamm, 2020), and a related strand of literature further investigates the heterogeneity of the parenthood effect on risk-avoidance by gender of the newborn (Pogrebna et al., 2018; Chew et al., 2018).

[2] The importance of children regarding parental old-age support has been documented by Banerjee et al. (2014), Imrohoroğlu and Zhao (2018) and Choukhmane et al. (2023). Specifically, both intergenerational transfers and parental old age support expectations measured based on population census and survey data document the importance of children in the support of elderly parents in China.

[3] We calculate the share of only children based on the 2015 census data. To the best of our knowledge, the 2015 census is the only census that asked adult individuals about their only-child status. In particular, the 2015 census asked whether married women aged between 15 and 50 and their spouses are only children or not. We then transform this woman-spouse-level information into individual-level information. Therefore, the share of only children captures the proportion of only children in the population of married women between 15-50 and their spouses.

ance roles due to shrinking family size and parental long-term care risks has been discussed in both public and academic contexts. In the arena of public discussions, for instance, the *Marriage and Family* magazine conducted a survey among parents with only one child regarding their perception of risk in their elderly support. As reported in the survey, 90% of the parents considered "*it is dangerous to have only one child for elderly support*", and 38% of only-child parents are specifically concerned about their children being ill or having accidents (Wang, 2004).[4] In scholarly discussions, Gui et al. (2004) and Mu (2009) raised similar concerns and discussed the greater risk of elderly support faced by only-child parents, as they have only one child to rely on, but their children may migrate or experience accidents that hurts the inter-generational support.[5] However, while existing discussions point to policy implications, empirical evidence is lacking to understand the relationship between the one-child family structure and parental risk behaviors in old age.

This paper examines the impact of having only one child on parental risk behaviors during the elderly stage. We use a nationally representative data set of people aged 45 and older from the China Health and Retirement Longitudinal Study (CHARLS) (Zhao et al., 2013). In terms of parental risk behaviors, we employ six measures to capture individual risk behavior in the health (Guiso and Paiella, 2004; Anderson and Mellor, 2008; Dohmen et al., 2011) and finance domains (Caballero, 1991; Fuchs-Schündeln and Schündeln, 2005; Dohmen et al., 2011). Specifically, these behaviors include (i) regular smoking; (ii) regular drinking; (iii) exercise habits (the number of days in a week with exercise for more than 10 minutes each day); (iv) whether there were household fitness expenditures in the last year; (v) whether the household has commercial insurance; and (vi) saving rate. Considering that family fertility decisions are made endogenously, we use the instrumental variable (IV) strategy based on the One-Child Policy in China as a quasi-experiment that affected the incidence of having only one child in the family.

Our instrumental variable for whether the family has only one child employs the variation of the One-Child Policy (OCP), which explicitly focuses on the margin of having one child. We construct a parent-level policy exposure variable to capture the impact of the policy using both regional and temporal variation. For regional variation, we examine the roll-out

[4] Recent news reports have also highlighted that the elderly support is more vulnerable in only-child families than in multiple-child families. For instance, the following news report (available at https://news.sina.com.cn/c/2006-09-13/ 135310999504.shtml) discusses various risk aspects faced by only-child families.

[5] In a similar vein, Naughton (2018) also discussed the potential impacts of China's recent demographic changes on household precautionary savings, pointing out that the family structure of one child can increase precautionary savings among the elderly.

of the Family Planning Commission (hereafter, the FPC) across prefectures, which was the leading local government organ in charge of the One-Child Policy enforcement since the 1980s. For cohort variation, we use a woman's year of birth to calculate her exposure to the region-specific enforcement of the One-Child Policy during her fertility-active window. Based on the two sources of variation, we then construct the parent-specific exposure to the policy enforcement, as measured by the share of affected years in the mother's fertility window weighted by the normalized age-specific fertility rate. The results show that the likelihood of having only one child increases significantly with mothers' policy exposure to the local establishment of the FPCs.

The main results indicate that parents who have more precarious old-age support with only one child show more risk avoidance behaviors, as *consistently* revealed in both health- and finance-related domains. Specifically, in health-related domains, only-child parents are less likely to smoke and drink regularly, spend more time on regular exercise, and are more likely to have family fitness expenditures. In finance-related domains, they are more likely to purchase commercial insurance and have a higher saving rate. The magnitudes of the effects are also economically sizable. Specifically, having only one child would decrease the probability of regular smoking and drinking by 45.5 and 80.7 percentage points, respectively, increase the number of days in a week with exercise over 10 minutes each day by 6.475, and increase the incidence of spending on fitness by 78.2 percentage points. In terms of finance-related behaviors, having only one child would increase the probability of holding commercial insurance by 20.3 percentage points, and increase the logarithm measure of the saving rate by 2.6, which is around a 1.24 standard deviation of the sample. Taken together, the estimated results indicate that only-child parents show more risk-aversion behaviors than multiple-child parents, and the effects are statistically and economically significant.

To examine the exclusion restriction condition of the instrumental variable strategy, we performed a set of falsification tests and placebo tests. In particular, we consider two alternative channels through which the enforcement of the OCP could affect parental risk outcomes other than by affecting their fertility outcomes, including (i) the enforcement of the policy can directly induce fertility-related conflicts and traumatic family experiences, which also affect parental risk behavior; and (ii) the OCP can also affect the structure of children's gender in addition to shrinking the number of children. By ruling out these alternative channels, the main results remain robust. Furthermore, we also conduct a placebo test by applying the analysis to the ethnic minority group, which was not systematically subjected to strict birth quotas under the OCP. The findings show no statistically significant relationship between the OCP and the risk behaviors of minority parents, which satisfies the exclusion restriction condition.

In addition to the main empirical analysis, we perform a series of robustness checks and heterogeneous analysis to the baseline IV analysis. Regarding robustness checks, the enriched analyses include controlling for province-by-cohort fixed effects, using alternative measures to calculate women's exposure to the policy, accounting for other fertility policy effects, and applying the inverse probability weighting (IPW) method to adjust for sample selection in the baseline analysis. By examining the robustness of the results across specifications, we find the effects of having only one child on parental risk behavioral outcomes remain robust and consistent with the baseline findings. Furthermore, we also examine the heterogeneity of the only child effects, specifically by gender of the first birth and by rural-urban status, respectively. The results show that only-daughter parents tend to be more risk-averse in certain behavioral domains. For rural-urban heterogeneity, we find that only-child parents in rural areas show more risk-avoidance behaviors than only-child parents in urban areas, and the differences are salient in both health-related and finance-related domains.

We also investigate the channels through which having an only child would affect parental risk behaviors with two components of analyses. In the first part, we examine the heterogeneity of only-child effects using both cultural and institutional variation regarding the importance of children in elderly support. The results show that the only-child risk-avoidance effects are intensified in regions with a stronger tradition of children serving as the main source of elderly support, while buffered in regions with a larger scale of civil servant systems that are associated with better institutional elderly support. In the second part of the analysis, we examine whether having one child can directly affect parental risk preferences. Our analysis employs individual-level data from China Family Panel Studies (CFPS) 2018, where respondents' risk preferences are elicited using a hypothetical lottery question that has been used in the literature (e.g., Hanaoka et al., 2018; Guiso and Paiella, 2004; Falk et al., 2018). The results show that only-child parents are more risk-averse than multiple-child parents in preferences. Taken together, the only-child parents not only show greater risk avoidance in terms of risk-related behaviors, but also show more risk aversion in terms of risk preferences.

Our paper contributes to three strands of literature. First, our findings add to the literature on the determinants of individual risk attitudes and risk-related behaviors, which currently include gender, age, and education (e.g., Borghans et al., 2008; Dohmen et al., 2011; Outreville, 2015; Schildberg-Hörisch, 2018), cognitive skills and psychological traits

(Dohmen et al., 2011; Mata et al., 2018), macroeconomic fluctuations (Malmendier and Nagel, 2011), health shocks (Decker and Schmitz, 2016), exposure to violence (e.g., Voors et al., 2012; Callen et al., 2014), near-miss accidents among automobile drivers (Shum and Xin, 2022), and natural disasters (e.g., Cameron and Shah, 2015; Hanaoka et al., 2018). Compared to the rich set of studies, understanding how family structure would affect risk attitudes and behaviors remains relatively scarce, especially considering the role of the family as a safety net in the life cycle. In this direction, current studies have examined two aspects of family structure, focusing on (i) the effect of parenthood and (ii) the number of brothers. Regarding the effect of parenthood, existing studies have shown that parenthood increases risk aversion around the time of the first childbirth (Görlitz and Tamm, 2020), increases the incidence of choosing safer jobs (DeLeire and Levy, 2004; Grazier and Sloane, 2008), increases riskier portfolio shares among young parents (Love, 2010), increases the willingness to pay for health-risk reductions (Cameron et al., 2010), and decreases male crime and tobacco and alcohol use (Kerr et al., 2011).[6] In addition to the parenthood effects, the literature also investigated the impact of having brothers on household risk-related financial outcomes. Specifically, Zhou (2014) shows that having more brothers can reduce the household savings rate, and Niu et al. (2020) shows that having more brothers increases the likelihood of participation in the stock market and the share of stocks in the household portfolio.

Given the existing scholarly work, our contribution to the literature is twofold. First, we study a new margin of family configuration – only-child families, by comparing only-child parents and multiple-child parents. This margin is especially important, as the family is essential for risk sharing and children play an important role in elderly support in China (e.g., Banerjee et al., 2014). Second, our analysis utilized variation from a quasi-experiment in China: the One-Child Policy. This policy explicitly shrinks the fertility margin from multiple children to just one child, which provides us with a valuable opportunity for identification. Using a combination of the policy variation by cohort and region, we find that parents having only one child as a result of the policy show greater behavioral risk avoidance in both the health and finance domains.

Our paper is also related to the literature on the consequences of the One-Child Policy (e.g., Zhang, 2017). This strand of literature examines the consequences of the OCP on children in the short- and mid-term by looking at the educational, marital, and labor market

[6] Existing studies have also examined the heterogeneous effects of parenthood by children's gender composition. For instance, Pogrebna et al. (2018) find that parents of daughters, unborn or recently born, become almost twice as risk-averse as parents of sons, and Chew et al. (2018) find that having sons significantly decreases parental risk aversion.

outcomes of the children, including their gender structure (e.g., Ebenstein, 2010), educational attainment, parental aspiration, and health outcomes (e.g., Liu, 2014; Li and Zhang, 2017; Qin et al., 2017), psychological traits and preferences (e.g., Cameron et al., 2013), marriage market outcomes (Huang et al., 2023), and labor market outcomes (e.g., Wang et al., 2017; Huang et al., 2021; Chen et al., 2023). Meanwhile, a few studies have shown the effects of the OCP on parental outcomes, including the impact of family size on maternal health (Wu and Li, 2012; Islam and Smyth, 2014).[7] However, investigating the long-term effects of the OCP on parents during their later years still remains scarce.

In this aspect, our paper adds to this literature by studying the impact of having only one child on parental outcomes during the elderly stage. Specifically, our paper looks at risk-related behaviors of elderly parents, motivated by the social and academic concerns in the tension between the deterioration of family insurance following the One-Child Policy and the long-term care risks faced by the elderly. Using both regional and temporal variation of the One-Child Policy, we find that only-child parents show more risk-avoidance behaviors in both health and financial domains, which is consistent with the qualitative discussions (e.g., Gui et al., 2004; Mu, 2009). Thus, our paper contributes to understanding the risk behavior of the elderly following fertility policies and speaks to both academic and policy interests.

Our paper also contributes to the literature that explains China's saving rate with demographic changes (e.g., Banerjee et al., 2014; Ge et al., 2018; Imrohoroğlu and Zhao, 2018; Choukhmane et al., 2023).[8] Specifically, this strand of literature is based on the tradition that children act as an important source of elderly support in Chinese society, and the main findings show that the OCP has contributed to the rise of the China's household saving rate. Compared to these studies, our paper adds to the literature by showing (i) the only-child parents' saving motives can be explained by *both* regional and cohort variation of the One-Child Policy; and (ii) parents also show more risk aversion in health-related behaviors, which is consistent with their risk behaviors in finance domains.

The remainder of this paper is organized as follows. Section 2 discusses background information, and Section 3 describes the data and measures of risk behaviors. Section 4

[7]Specifically, Wu and Li (2012) find that mothers with fewer children are in better health in terms of calorie intake, BMI, and blood pressure; and Islam and Smyth (2014) provide complementary findings in the long term.

[8]Scholars have also examined the impact of sex ratios on saving rate. For instance, Wei and Zhang (2011) provides a competitive saving motive at the household level by demonstrating that families with boys facing intensified marriage market competition save more in response to biased sex ratios. In line with this perspective, our study examines an additional household-level explanation for saving rates as discussed by Naughton (2018), that is, only-child parents face greater risk in elderly support and conduct more precautionary savings.

presents the identification strategy, and Section 5 reports the main findings. Section 6 shows the heterogeneity analysis and Section 7 discusses the mechanisms. Section 8 concludes.

# 2 The One-Child Policy and Institutional Reforms

## 2.1 The One-Child Policy and Shifting Family Structure

The Family Planning Policies in contemporary China can be traced back to the fertility policy campaigns during the 1950s-1960s, which aimed to alleviate the socioeconomic and environmental burden of a burgeoning population (Banister, 1987; Peng, 1997; Yang, 2004). During the 1960s, the fertility policy was featured by relatively mild implementation, promoting late marriages and introducing family planning technology. During the 1970s, the fertility policy was tightened following the introduction of the "Later, Longer, Fewer" policies. This policy package enhanced fertility-related restrictions in various aspects, including later marriages, longer birth spacing, and fewer children at the core. As evaluated by recent studies, this policy package during the 1970s played an important role in the fertility decline since the early 1970s (Chen and Huang, 2020; Chen and Fang, 2021). Thereafter, the One-Child Policy was proposed in 1979 and propagandized in 1980 at the national level, noted by "the Open Letter of the Central Committee of the China Communist Party to All the Party Members and Youth League Members" (Yang, 2004). With rich relaxation and variation, the One-Child Policy was formally replaced by the Two-Child Policy until January 2016.

Compared to the early phases of the fertility policies during the 1960s and 1970s, the One-Child Policy was featured, predominantly restricting ethnic Han Chinese couples to only one child during their lifetime. At the same time, only-child families — those parents who have only one child during their lifetime — have been a rising family configuration in recent decades. As shown in Figure 1, along with the declining total fertility rate in the past decades, the One-Child Policy led to a noticeable change in the structure of the household in China. Specifically, we employ data from the China census 2015 to illustrate the trend of only children in recent years. As shown by the trends, the share of only children increased by birth cohort from 2% in 1965 to approximately 10% in 1990. The calculated pattern is consistent with the demographic prediction conducted by Wang (2009), that the stock of only children has increased over the past decades.

## 2.2 The One-Child Policy and Institutional Reforms

The institutional setup has played a crucial role in the implementation of fertility policies. For each stage of the fertility policies in contemporary China, the governmental organizational setting varied.

Figure 2 summarizes the evolution of fertility policies and the corresponding government organizational settings. In the left panel, we illustrate three main phases of the fertility policies following Yang (2004), including the early fertility campaigns during the 1950s and 1960s, the "Later, Longer, Fewer" campaigns during the 1970s, and the One-Child Policy from 1979 to 2016. In the right panel, we show the evolution of government organizations responsible for implementing local fertility policies (Huang, 2014; Chen, 2015; Zhang, 2015). In Figure 2, we illustrate that the Family Planning Commission was responsible for implementing the OCP, as a powerful bureau-level government organ parallel to the Health Bureau. In this paper, we explicitly examine the timing of the FPC establishment by localities so as to study the policy effects from an organizational perspective.[9]

As shown in Figure 2, during the early phases of the fertility policies, the Family Planning Office and the Family Planning Leading Group were the main government organizations responsible for their implementation during the 1960s and 1970s (see Chen and Huang (2020); Chen and Fang (2021) for discussions on the Family Planning Leading Group). In terms of organizational structure, both the Family Planning Office and the Family Planning Leading Group work as subordinates under the Health Bureau. After the announcement of the One-Child Policy, organizational reforms were initiated to enhance policy implementation. At the local level, the FPC was established as an empowered unit separated from the Health Bureau and elevated to the bureau level in administrative ranking. Subsequently, the FPC became the predominant government institution responsible for policy implementation and was directly linked to the rigid implementation of fertility policies, such as assigning birth quotas and imposing penalties for excessive births in localities (Zhang, 2015). The specific duties of the FPC include: (i) specifying and implementing the regional fertility policy targets; (ii) supervising policy implementation among the lower-level government units; (iii) promoting policy propaganda; (iv) safeguarding the allocation of contraceptives and encouraging family planning studies; (v) supervising financial plans; (vi) cultivating teams of cadres; and (vii) organizing international events on family planning (Peng, 1997; Yang, 2004). Thus,

[9]Previous studies have discussed two important aspects of the organizational structure of fertility policy implementation. For one, the literature on the "Later, Longer, Fewer" policy has demonstrated the impacts of the provincial Family Planning Leading Group on aggregate fertility outcomes and parental outcomes (Chen and Huang, 2020; Chen and Fang, 2021). For the other, existing studies regarding the One-Child Policy have also examined the effect of the provincial roll-out of other policy-related organizations (e.g., Edlund et al., 2007; Jia and Persson, 2020; Chen et al., 2023), including the Family planning science and technology research institute, Family planning education center, Family planning association. These organizations were under the leadership of the FPC (Peng, 1997).

the establishment of the FPC across regions provides variation regarding local-level policy enforcement.

To proceed with our analysis of the effects of local FPC establishment, we collected the FPC establishment year at the prefecture level from a battery of sources, including prefecture yearbooks, local gazetteers, the PKU-Law dataset, and the Baidu encyclopedia. Among the 126 prefectures covered by CHARLS data, we found the accurate FPC establishment year for 83 prefectures. For the rest of the 43 prefectures, we do not have prefecture-level information for the precise establishment year of the FPC from the above available sources, and we use the province-level FPC year as the closest proxy for these 43 localities after collecting the original information from Peng (1997). In Figure 3, we show the distribution of the FPC establishment years. As shown in the figure, the establishment year of FPC varied widely across a period of more than seven years, and in around 60% of the prefecture, FPCs were established in 1983.[10] Moreover, there exists considerable variation in the prefecture-level FPC establishment within province.[11]

To illustrate the evolution of fertility-related organizations, we examine an example from Nanchang prefecture in Jiangxi province. As shown in Appendix Figure A1, the history of the family planning organizations in Nanchang underwent major changes. During the earlier stage of the fertility policies before the OCP, the Family Planning Office and the Leading Group were the dominant organizations responsible for implementing the relevant policies. Since February 1982, the Nanchang government elevated the Nanchang FPC to the bureau level, which stands as an organ parallel to the health department (Nanchang Gazetteer Compilation Committee, 1997). The Nanchang FPC then acted as the predominant institution responsible for implementing the OCP afterward.

# 3 Data and Measurement

## 3.1 Data and Sample

Our data come from the baseline wave of the China Health and Retirement Longitudinal Study (CHARLS) that was conducted in 2011, covering 28 provinces, 150 counties, and

[10] While most local FPCs were established after the foundation of the national FPC, there are two exceptions when local FPCs were established in earlier years. Specifically, Shangrao prefecture established its FPC in 1979, as documented by Xiao (1995). The other exception is the case of Tianjin, where the local FPC was established in October 1978, as documented by Zhu (1991). In this analysis, we use the original information of the FPC establishment years for these two localities, as documented by the sources.

[11] By doing a regression using the prefecture-level data, we find around 50% of the variation of the FPC establishment year can be explained by the province fixed effects.

450 townships or streets (Zhao et al., 2013). CHARLS was designed to understand the social and economic status of the Chinese elderly over 45 years of age. It is comparable to other international studies targeting the elderly, including the Health and Retirement Study (HRS) in the U.S., the English Longitudinal Study of Aging (ELSA) in the UK, and the Survey of Health, Aging, and Retirement (SHARE) in Europe. The CHARLS consists of a representative sample of the Chinese population aged 45 and older using the Probability Proportional to Size (PPS) sampling method and collects rich information at the individual, household, and community levels. For our study, the CHARLS provides detailed information on the elderly and their children, which is essential for the empirical analysis.

As the focus of our study is to understand whether having just one child would affect parents' behavioral outcomes, we apply several sample restrictions to the raw data set of CHARLS 2011, which includes 17,780 individuals aged 45 and older in 10,257 households. First, we restrict our sample to people of Han ethnicity, born from 1935 to 1966, which includes a group whose main fertility-active windows cover both pre- and post-periods of the One-Child Policy. We then restrict the sample to those who have married only once and have at least one child, as the inclusion of people who are in their second or consecutive marriages might induce noise when measuring their exposure to fertility policies. Lastly, we keep parents with sufficient non-missing information on a battery of basic demographic and socio-economic variables, including gender, age, household registration (*Hukou* status), educational attainment, spousal age, spousal education, widowhood status, and smoking and drinking history.

## 3.2 Measurement of Risk-related Behaviors

CHARLS provides us with a rich set of information regarding individual risk-related behavior. In terms of parental risk behaviors, we employ six measures to capture individual risk behavior in the health (Guiso and Paiella, 2004; Anderson and Mellor, 2008; Dohmen et al., 2011) and finance domains (Caballero, 1991; Fuchs-Schündeln and Schündeln, 2005; Dohmen et al., 2011). The behaviors include (i) regular smoking; (ii) regular drinking; (iii) exercise habits, measured as the number of days in a week with exercise for more than 10 minutes each day; (iv) whether there were any household fitness expenditure in the last year; (v) whether the household has commercial insurance; and (vi) saving rate. Specifically, the regular smoking variable is a dummy variable that indicates whether the respondent currently has a habit of smoking, and the regular drinking variable is a dummy variable that indicates whether the respondent drinks more than once a month. As such, both smoking and drinking capture habits rather than occasional behaviors. The exercise variable encompasses moderate physical activities specified in the CHARLS questionnaire, such as brisk walking, practicing Tai Chi, and regular cycling. We construct a measure denoting the number of days in a week the respondent engaged in these exercises for more than 10 minutes on a daily basis.[12] The fitness expenditure variable is a binary indicator of whether the household had any fitness-related expenditure in the last year, covering items such as fitness equipment and health products. The commercial insurance variable is a dummy variable at the household level that indicates that the husband or wife has commercial insurance. In terms of saving rate, we use two measures following the literature. The first measure follows the classic definition of saving rate (e.g., as in Imrohoroğlu and Zhao, 2018; Choukhmane et al., 2023), which is calculated as $saving\ rate_1 = 1 - expenditure/income$. Given possible noise around the tails, we restrict the sample by keeping its value between -3 and 1, which accounts for 85.8% of the observations.[13] For robustness, we employ a second measure following Wei and Zhang (2011), where the saving rate is measured as $saving\ rate_2 = ln(income/expenditure)$. Together, the battery of various measures can capture individual risk-related behaviors on both the health and finance dimensions.

Based on these measures, we further construct a summary index of risk-related behaviors, using the approach of Kling et al. (2007) in two steps. First, we harmonize the directions of all the above risk behavior variables, so that a higher value of the index captures a greater tendency of risk avoidance. Specifically, among the various components, a lower incidence of regular smoking and drinking, more days in a week conducting exercise, a higher incidence of fitness expenditures, the holding of commercial insurance, and higher saving rates indicate a greater tendency of risk avoidance. The summary index variable is then generated as the average of the standardized values of all risk behavioral variables, and we label this variable as $Risk\ Index$ 1. For robustness check, we also calculate a second risk behavioral index excluding the exercise variable, as only a randomly selected group of respondents was asked to answer the question in the survey; and we label this variable as $Risk\ Index$ 2. Table 1 presents the summary statistics of key variables in the sample.

As shown by Table 1, our main sample consists of 13,068 individual parents, whose ages range from 45 to 76. In the sample, 18% of the parents have only one child. Around 21% of the parents are affiliated with urban *Hukou*. In terms of educational composition, about 42% of parents are illiterate or semi-illiterate, and those who have an educational

[12] Notably, only a randomly selected group of respondents was asked to answer this question in the survey.

[13] Truncated values are often used in studies on saving rates to deal with extreme values (e.g., Zhou, 2014; Chen et al., 2019). Our current truncated value is set to -3 and 1, and the main results are robust to alternative choices of truncation values.

attainment of primary, middle, and high school education or above account for 22%, 22%, and 13% of the sample, respectively. On average, the age gap between couples is 2.16 years, and husbands are better educated than wives. For risk-related behaviors, around 33% and 26% of the parents currently have smoking and drinking habits, they on average have 3.62 days in a week conducting excise over 10 minutes a day, 24% of the parents spent money on fitness in the previous year, and 5% of the parents have commercial insurance. Regarding the saving rates, the average saving rate using the first measure ($saving\ rate\ 1 = 1 - expenditure/income$) is 50%, and the average saving rate using the second measure ($saving\ rate\ 2 = ln(income/expenditure)$) is 1.2. For the summary index of risk-related behavior, the average value of $Risk\ index\ 1$ (the calculation includes all behavioral variables) is 0.034, and the average value of the second risk index $Risk\ index\ 2$ (the calculation excludes the exercise variable) is similar, which is 0.033.

# 4 Empirical Strategy

In this section, we estimate the effect of having an only child on parents' risk behavioral outcomes. Our baseline estimation equation is as follows:

$$Y_{ijc} = \alpha + \beta Only_{ijc} + X_i'\Gamma + \delta_j + \theta_{kc} + T_{jc} + \varepsilon_{ijc} \tag{1}$$

Here $i$ indexes parent, $c$ represents the parent's year of birth, $j$ indicates prefecture, and $k$ indicates the four major macroeconomic regions in China.[14] $Y_{ijc}$ is the outcome variable that captures parental risk behaviors. $Only_{ijc}$ is the key explanatory variable, which is a dummy variable that indicates whether the parent has only one child. The vector $X_i'$ includes a battery of control variables, including gender, educational attainment, *Hukou* status, the indicator for widowhood status, parental spousal age gap, spousal educational gap, and baseline risk behaviors before having the first child.[15] In addition to the above covariates, $\delta_j$ represents the prefecture fixed effects; $\theta_{kc}$ represents the region-by-cohort fixed effects; and $T_{jc}$ represents the prefecture-specific cohort linear trends. $\epsilon_{ijc}$ is the error term and we use robust standard errors clustered at the prefecture level. The coefficient of interest is $\beta$, which captures the difference in the risk behavioral outcomes between the only-child parents

[14] Here the regions are classified following the four major economic regions of China by definition of the National Bureau of Statistics, namely the Eastern, Western, Central, and Northeastern region.

[15] Here we included the baseline risk behavior variable, considering that the only-child parents may differ in risk behavior even before having the first birth. Specifically, we use information on individual drinking and smoking history to capture individual baseline risk behavior, using a dummy variable that indicates whether the respondent had a habit of smoking or drinking before the first child was born.

and multiple-child parents.

When estimating the effect of having an only child on parental risk behavior outcomes using Equation (1), the ordinary least squares (OLS) method is informative in revealing the correlation between the only-child family type indicator and parental risk behaviors. However, the OLS estimation may not deliver the causal effect due to several identification challenges. In particular, parents may endogenously decide on the number of children, which leads to omitted variable bias, and there could be unobserved parental characteristics that affect their fertility decisions and risk behaviors at the same time. In addition to omitted variable bias, we may also face the endogeneity issue arising from measurement error. As demographic studies document (e.g., Cui et al., 2013), misreporting is often an issue for the newborn in the population census. However, since the CHARLS survey targets the elderly population above age 45 who do not face systematic incentives to misreport newborn children, we consider misreporting less likely to result in measurement error issues.

To meet these empirical challenges, we use the instrumental variable approach using the variation generated by the implementation of the One-Child Policy in China. The policy has several merits in facilitating our analysis. First, it explicitly focused on the margin of having one child. Second, as discussed in Section 2, its implementation had regional and temporal variations, as captured by the establishment of the Family Planning Commission, which provides variation for identification. Specifically, our instrumental variable for only child status combines two sources of variation, by comparing mothers of different cohorts and prefectures that established the FPC in different years.

Using the above information, we then construct a measure of mother-level exposure to the Family Planning Commission. The idea is that the implementation of fertility policies affects women's fertility outcomes by imposing restrictions on women's fertility ages. This strategy is conceptually in line with Wu and Li (2012), which used 1979 as the cutoff year for the One-Child Policy and calculated mothers' exposure to the OCP by cohort, and that of Chen and Huang (2020) and Chen and Fang (2021), which examined the effect of the Later-Longer-Fewer (LLF) fertility policies in the 1970s China and calculated mothers' exposure to the LLF using the year of establishment of the Family Planning Leading Group by province. Compared to these existing measures, our measure is based on a new set of variations driven by OCP implementation across regions – the establishment of the Family Planning Commissions, thus enriching the regional policy variation compared to Wu and Li (2012), and extending the horizon of fertility policy implementation in addition to Chen and Huang (2020) and Chen and Fang (2021).

Formally, our definition of a mother's policy exposure to the FPC of the prefecture is:

$$MFPC_{jc} = \sum_{a=15}^{49} [ASFR(a) \times I[a + c \geq fpc\_year_j]]/CFR \tag{2}$$

Being consistent with the baseline notations, here $j$ indicates the prefecture and $c$ represents the year of birth of the mother. The parameter $a$ ranges from 15 to 49, indicating the fertility-active ages following the definition of the fertility-active ages of women by the China's National Bureau of Statistics. $ASFR(a)$ is the age-specific fertility rate (ASFR) of women of age $a$. The ASFR in our main analysis is calculated based on CHARLS 2011, using the fertility histories of women born between 1925 and 1940, which constitutes the group whose main fertility-active windows are before 1979. We used the completed fertility rate (CFR) of the same group of women to normalize the scale from 0 to 1, which is also calculated based on CHARLS 2011.[16] In addition to the ASFR distribution calculated from CHARLS, we also perform a robustness check using alternative sources of ASFR from Coale and Chen (1987). $I[a+c \geq fpc\_year_j]$ is an indicator variable that captures whether a woman would be exposed to the FPC in her fertility window, where $fpc\_year_j$ indicates the year of establishment of the FPC in prefecture $j$.[17] Intuitively, the $MFPC_{jc}$ captures the intensity of a woman's exposure to the policy implementation during the fertility-active window, with both regional and cohort variations.

Following the definition, we further use an example to illustrate the variation of the MFPC. In Figure 4, we consider a woman born in 1950 who lives in Shanghai. Shanghai's FPC was established in 1982, and the fertility ages of the woman would be affected during her 32 to 49. Therefore, the mother's policy exposure (MFPC) is the share of FPC-affected ages during her whole fertility window weighted by ASFR, which is graphically captured by the proportion of the grey area below the red line in Figure 4. Intuitively, we would expect a positive correlation between a mother's exposure and fertility outcomes.

To examine the performance of the first stage, we study the correlation between a mother's policy exposure and fertility outcomes. Figure 5 shows two summary trends: (i) the average MFPC by mothers' birth cohort, and (ii) the share of women with only one child

[16] For details on the calculation of the ASFR and CFR, please refer to Appendix A.

[17] In doing so, we assume the mothers resided in the same prefecture at the time of the survey as the prefecture in which they gave birth. In the data, 97.7% of the parents live in the same county/district where the first child was born. The low migration rate may be due to the fact that the respondents are relatively old (the average age of our sample is 57.9) and migration in China was still quite restrictive when they were young.

by cohort. We observe that the share of only-child women started to increase roughly around the 1950s cohorts. Specifically, the share of women with only one child is 8.5 percent among the pre-1950 cohort, and the share increased to around 30 percent among the youngest cohort in the sample (i.e., those born in 1966). Tracking the average policy exposure of women (the MFPC), we find that those women who are of younger cohorts are on average more affected than those of older cohorts. In summary, the general trends regarding women's policy exposure and the share of only children increased hand in hand. Formally, we present the first-stage regression equation:

$$Only_{ijc} = \alpha_1 + \beta_1 MFPC_{jc} + X_i'\Gamma_1 + \delta_j + \theta_{kc} + T_{jc} + u_{ijc} \tag{3}$$

As before, $i$ indexes parents, $j$ indicates prefectures, $c$ represents the year of birth of parents, and $k$ indicates the major macroeconomic regions. $Only_{ijc}$ is a dummy variable that indicates whether the parent has only one child. Since $MFPC_{jc}$ is calculated based on the year of birth of the woman and the prefecture of residence, for men, it is calculated based on the birth cohort of their wives and the prefecture of residence. The vector $X_i'$ denotes a battery of control variables of the parent, including gender, educational attainment, *Hukou* status, a widowhood indicator, the spousal age gap, the spousal educational gap, and smoking and drinking history. $\delta_j$ and $\theta_{kc}$ represent the prefecture fixed effects and the region-by-cohort fixed effects, respectively. $T_{jc}$ represents prefecture-specific cohort linear trends. $u_{ijc}$ is the error term, clustered at the prefecture level. $\beta_1$ captures the effect of women's policy exposure on the incidence of having only one child in the family. This empirical strategy is able to account for all time-invariant differences across prefectures by prefecture fixed effects ($\delta_j$), and all macroregion-by-cohort characteristics observed and unobserved by the region-cohort fixed effects ($\theta_{kc}$). Furthermore, we control for prefecture-specific cohort linear trends to account for differential linear trends in outcome variables in all prefectures ($T_{jc}$). Therefore, our IV estimate exploits the variation in the status of only children by comparing mothers from different cohorts and prefectures that established the FPC in different years.

Next, we discuss the validity of the IV. The validity of the IV strategy relies on the condition that the instrumental variable is correlated with the only-child status, and it does not affect parental risk-related outcomes through additional channels. While the correlation condition can be directly verified by the first-stage results, the exclusion restriction requirement is more subtle to examine. In particular, one might wonder whether the establishment of the FPC would be correlated with local socioeconomic characteristics. To investigate this issue, we examine a simple correlation analysis between the timing of the FPC setup and the local socioeconomic conditions prior to the OCP in Table 2.

In the analyses, we employ two sets of information to capture the local socioeconomic conditions prior to the OCP: (i) prefecture-level characteristics, including educational, occupational and employment structure of the population aged 25-50 in 1977, and completed fertility rate among women 45-54 in 1977, which are calculated using the 1982 Population Census;[18] and (ii) province-level characteristics, including provincial GDP per capita in 1977, crude birth rate in 1977, and urbanization rate in 1977, which are from the National Bureau of Statistics (2010). In Columns (1) to (4), we progressively control for prefecture-level characteristics, provincial-level characteristics, and provincial fixed effects. The results show that there is no statistically significant relationship between the FPC establishment year and the regional characteristics. In Columns (5) to (8), we use the order of establishment among the prefectures as an alternative outcome variable. In parallel, we do not find evidence for the local socioeconomic conditions explaining the relative timing of the FPC establishment. Taken together, this set of analyses shows that the years of FPC establishment are not correlated with the main regional economic variables prior to the policy, and we provide additional evidence and robustness checks in Section 5.3 to assess the validity of the exclusion condition.

# 5 Main Results

## 5.1 The OLS Estimates

We first discuss the OLS estimates of Equation (1). In Table 3, we examine the correlation between having only child and parental risk behavioral outcomes. Columns (1) to (4) present the health-related risk behavioral outcomes, including regular smoking, regular drinking, the number of days with exercise for more than 10 minutes per day during the week, and whether the household had any fitness expenditure in the last year. In Columns (5), (6), and (7), we examine the commercial insurance holding and saving rates (with two measures). In Columns (8) and (9), we employ the risk indices to capture the general tendency of risk avoidance. Regarding the control variables, we employ two different sets of control variables in Table 3, in Panel A and Panel B. Specifically, Panel A includes the individual-level control variables,

[18]Specifically, the prefecture-level socioeconomic variables include: the share of people with junior high school education among the population aged 25-50 in 1977, the share of people with senior high school education and above among the population aged 25-50 in 1977, the share of secondary workers among the population aged 25-50 in 1977, the share of tertiary workers among the population aged 25-50 in 1977, the unemployment rate of the population aged 25-50 in 1977, and the completed fertility rate of women aged 45-54 in 1977.

which would partially account for the potential confounders that may be correlated with both parental risk behavioral outcomes and having only one child. Panel B further adds prefecture fixed effects, macroeconomic region-cohort fixed effects, and prefecture linear trends to those controls in Panel A.

As shown by the results of the two panels, having only one child is weakly correlated with parental risk behavior. In particular, the results in Panel A show that only-child parents are more likely to drink regularly, more likely to have fitness expenditures, spend more time on exercise during the week, are more likely to have commercial insurance, and they also have higher saving rate. On the other hand, in Panel B, when we account for a richer set of controls, having only one child is not correlated with parental risk behavior. Together, the OLS results do not reveal consistent patterns with respect to the relationship between having one child and parental risk behaviors. At the same time, the interpretation of the OLS results is unclear, given the empirical challenge of omitted variable bias that we discussed. Next, we move on to the instrumental variable analysis following the setting of Section 4.

## 5.2 The IV Estimates

Table 4 reports the IV estimates of the effects of having only one child on parental behavioral outcomes. In this table, Panel A presents the first-stage estimates, Panel B reports the second-stage estimates, and Panel C shows the reduced-form estimates. The IV estimation controls for our baseline set of control variables, as described in Section 4.

In Panel A, the first-stage estimates show that a mother's exposure to the establishment of the Family Planning Commission during her fertility-active window positively predicts the incidence of having only one child. In terms of the magnitude, taking Column (1) as an example, the results show that increasing a women's exposure from zero to one (i.e., comparing a woman whose fertility-active window is completely covered by the FPC establishment and a woman whose fertility-active window without any overlap with the FPC establishment) would increase the probability of having only one child by 20.4 percentage points. This is considerable in terms of the economic effect, which is around 53.1% (=20.4/38.4) of one standard deviation of having only one child in the sample. Furthermore, the first-stage KP $F$-statistics are beyond the heuristic standard to detect weak instrumental variables (Stock and Yogo, 2005) through different columns of outcomes with varying sample sizes. Together, the first stage results show that the establishment of the FPC has affected fertility outcomes by increasing the incidence of having only one child under the One-Child Policy.

We then move on to discuss the second-stage estimates. As shown in Panel B, the estimates indicate significant effects of having only one child on parental risk-related behavioral outcomes. Specifically, only-child parents show more risk avoidance in health and finance-related behaviors, by being less likely to smoke and drink regularly, spending more time on regular exercise, showing a higher likelihood of having fitness expenditures, being more likely to have commercial insurance, and having a higher saving rate. The results consistently reveal that only-child parents are more risk-averse in both the health and finance behavioral domains. In terms of magnitude, the effects are also economically significant and meaningful. Specifically, for health-related behaviors, having only one child would decrease the probability of regular smoking and drinking by 45.5 and 80.7 percentage points, respectively, increase the number of days in a week with exercise over 10 minutes a day by 6.475, and increase the incidence of fitness expenditure by 78.2 percentage points.[19] For finance-related behaviors, having only one child would increase the probability of commercial insurance holding by 20.3 percentage points. For saving rate, having only one child would increase the share measure of saving rate (i.e., $saving\ rate\ 1 = 1 - expenditure/income$) by 89.1%, and the logarithm measure of saving rate (i.e., $saving\ rate\ 2 = ln(income/expenditure)$) by 2.6, which is around 1.24 standard deviation of the sample. We observe that the IV estimates are considerably larger than the OLS estimates, which suggests that selection into an only-child family biases the OLS estimates. For example, parents who anticipated more generous institutional elderly support may be more willing to have just one child, and they may also show less risk avoidance in behaviors. This will result in an under-estimation of the only-child effects in the OLS analysis. At the same time, it should be noted that IV estimates represent the local average treatment effect (LATE), which is the average treatment effect for compliers. In our setting, the compliers are those parents who would have only one child due to the enforcement of the One-Child Policy. We are interested in this group because the compilers of the policy are directly relevant to policy evaluation and welfare implications.

Panel C examines the reduced form results, which reveal the relationship between mother's exposure to the FPC and parental risk-related behavioral results. The correlation is consistent with the first- and second-stage IV estimates: those parents whose fertility windows are more affected by the FPC show more risk-avoidance behaviors in health- and finance-related domains.

[19] We consider an alternative explanation for the results of exercise, that parents with only one child may have more free time with fewer obligations for taking care of grandchildren, thus being more likely to conduct weekly exercises. To examine this hypothesis, we study the effects of having only one child on time-using behaviors. As shown in Table A1, we find that having only one child is not significantly related to weekly hours of taking care of grandchildren, hours of sleep at night, and minutes of nap during the noon.

### 5.3 Examination of the Validity of the Exclusion Restriction

The IV estimation above rests on the assumption that mothers' exposure to the FPC would only affect parental risk-related behaviors by affecting their fertility outcomes. In this section, we conduct a set of falsification tests to examine the possible alternatives through which the enforcement of the OCP could affect parental risk-related outcomes other than affecting their fertility outcomes. In particular, we consider two alternative channels, including (i) the enforcement of the policy may directly induce fertility-related conflicts and traumatic family experiences, which also affect parental risk behavior; and (ii) the OCP may also affect children's gender structure in addition to shrinking the number of children.

We start by evaluating the first alternative explanation through the conflict channel. As discussed by existing scholarly work (e.g., Greenhalgh and Winckler, 2005), the enforcement of the OCP has often been accompanied by mandatory family planning surgeries and abortions. These traumatic experiences may also likely have affected individual risk attitudes and behaviors.[20] To examine this alternative channel, we conduct two robustness checks.

In Table 5, we evaluate this alternative channel by accounting for the traumatic experiences of families, where the traumatic experience is proxied by whether they have experienced an induced abortion. As shown, the results after controlling for this proxy remain robust, which is consistent with the baseline findings. Meanwhile, we find that the experience of induced abortion is associated with a greater degree of risk-taking in health-related behaviors; and we do not find a statistically significant relationship between the experiences and finance-related behaviors. As such, it seems that traumatic experiences do not consistently predict parental risk-related behavioral outcomes in different domains.

In addition to directly accounting for individual experiences, we conduct another analysis to evaluate the conflict channel. In particular, we follow the findings of Peng (2010), which documented that the enforcement of the OCP and the induced conflicts could have been mitigated by the presence of clans. Therefore, if the conflict is the main channel through which OCP enforcement would affect parental risk-related behavior, we would expect the reduced-form association between mothers' exposure to the FPC and risk behavioral outcomes to be weaker in regions with more powerful clans, where parents were less likely to witness and experience conflicts induced by the OCP.

[20] However, there is little consensus about the direction in which such negative shocks affect risk attitudes and behaviors. In extant studies regarding conflicts and traumatic experiences as the determinants of risk preferences, the findings are mixed, showing that conflicts and traumatic experiences could increase (e.g., Callen et al., 2014; Kim and Lee, 2014) or reduce (e.g., Voors et al., 2012) the tendency of risk aversion in different contexts.

In this vein, we conduct a heterogeneity analysis on the effect of mothers' exposure to the FPC by the local strength of the clans. We add an interaction term to the reduced-form analysis, namely the interaction between MFPC and the local strength of the clans. To proxy the strength of local clans, we use the density of genealogies at the prefecture level and divide the prefectures into high and low density of genealogies groups by the median value, where the variable *Low Clan Strength* indicates whether the number of genealogies per capita is below the median at the prefecture level.[21] Thus, the effects of mothers' exposure to the FPC in regions with a powerful clan would be captured by MFPC in our regressions, and the heterogeneous effects between regions with different clan strengths are captured by the interaction term (i.e., *Low Clan Strength* $\times$ MFPC).

As shown by the results presented in Table A2 in the Appendix, we find that, consistent with the baseline results in Panel C of Table 4, the main effects of our interest remain robust that mothers' exposure to the FPC affected parental risk-related behaviors. Moreover, the interaction effects are mostly not statistically significant. The results thus provide another complementary evidence that local conflicts are not likely to be a main channel through which OCP can affect parental risk behaviors.

After evaluating the conflict aspect, we then move on to examine another channel, that the One-Child Policy may also alter children's gender structure, which could affect parental risk-related behaviors. This is an important alternative explanation as well, since the OCP had contributed to higher sex ratios (e.g., Ebenstein, 2010), and Wei and Zhang (2011) have documented that biased sex ratios increased household saving rates, as parents with sons faced intensified marriage market competition financially.[22] In this scenario, the positive correlation between mothers' exposure to the FPC and higher saving rates would capture the effect of the gender structure of children within the household on saving rate.

To account for this alternative channel, we directly control for the gender and marriage structure of the children within the household. In particular, our control variables include two indicator variables: (i) whether the parents have unmarried sons; and (ii) whether the parents have unmarried daughters. Following Wei and Zhang (2011), the theoretical prior would be that unmarried sons would predict a higher saving rate. As shown in Table A3 in the Appendix, we find that controlling for children's gender and marital status does not alter the main findings of the IV estimates for the only-child effects on parental risk outcomes.

[21] The genealogy data is from Wang (2008), which documented a catalog of genealogies.

[22] In a similar vein, Cameron et al. (2019) finds that China's high sex ratios resulted in greater financial pressure due to the competitive marriage market, and the high sex ratios are associated with greater risk-taking amongst males, which partly explains the increase in criminality.

In addition, consistent with the prediction of Wei and Zhang (2011), we do find that having unmarried sons predicts a higher saving rate within the household while having unmarried daughters is not statistically correlated with the household saving rate. Taken together, the findings further support the main findings by addressing specific concerns regarding the exclusion restriction.

In addition to ruling out alternative channels, we also conduct a placebo test to provide further support for the validity of the exclusion restriction assumption. The idea of the placebo test is related to the variation of the OCP by ethnicity. According to policy rules, ethnic minorities were legally allowed to have two or more births or were not even subject to the OCP (Li et al., 2011; Huang et al., 2023). If the above exclusion restriction is a valid assumption, then the exposure of ethnic minority mothers to the FPC should not directly affect their risk-related outcomes. By contrast, if there are significant correlations between ethnic minority mothers' exposure to the FPC and their risk behaviors, it is likely there could be other channels through which the IV would affect the outcomes.

In Table 6, we present the results of this analysis, where a reduced-form analysis is conducted by examining the effect of ethnic minority's exposure to the FPC on their risk-related outcomes.[23] Panel A restricts the sample to individuals who are ethnic minorities, whereas Panel B restricts the sample to individuals who themselves or their spouses are ethnic minorities. Through Columns (1) to (9) of both Panels, we find the main estimates are statistically insignificance in most cases, and the direction of the effects is mixed across various risk-related outcomes. The placebo results thus again support the validity of the exclusion restriction condition.

## 5.4 Robustness Checks

In this section, we undertake a set of robustness checks. The first robustness check applies province-by-cohort fixed effects compared to the macroregion-by-cohort fixed effects in our baseline regression. This set of controls accounts for all province-by-cohort characteristics observed and unobserved as a more stringent set of controls. As shown by Panel A of Table 7, the IV estimates of the effects of having only one child on parental risk behavioral outcomes remain robust and consistent with the baseline findings.

The second robustness check that we conduct is using an alternative ASFR distribution

[23]Since the policy exemption term that allowed excess birth quotas were only assigned to national minorities with less than 10 million members in some provinces (Scharping, 2003), we treat parents belonging to the Zhuang ethnic group (with a national population of over 13 million according to the 1982 Population Census) as a non-ethnic minority and exclude them from the sample we use in this part.

when calculating women's exposure to the FPC. In our baseline analysis, the ASFR distribution is calculated based on the CHARLS 2011, among the group of women born between 1925 and 1940. In the robustness check, we employ an alternative distribution from Coale and Chen (1987), where the ASFR distribution was in 1977. As the source does not include complete fertility rate (CFR), we use total fertility rate (TFR) in 1977 as the denominator of the exposure, namely:

$$MFPC_{jc} = \sum_{a=15}^{49}[ASFR_{1977}(a) \times I[a + c \geq fpc\_year_j]]/TFR_{1977} \tag{4}$$

Using this alternative calculation of the MFPC, we present the IV estimates in Panel B of Table 7. As shown, the second-stage results are largely comparable to those from the baseline analysis. The first-stage results are a bit weaker, as suggested by the $F$-statistics, yet they are also above the heuristic standards for weak instrumental variables in most cases.[24]

The third robustness check we conduct is applying the inverse probability weighting (IPW) method to adjust for sample selection of the baseline analysis. We obtain original sampling weights that correct for the non-response individuals from the dataset and compute the probability that a parent in the original sample is included in our baseline sample.[25] The final sampling weights are derived by multiplying the original weights by the inverse of the computed probability and then truncated at the $99.5^{th}$ percentile to ensure that no observation is unduly weighted. The IV estimates applying the sampling weights are presented in Panel C of Table 7. The results also show that the estimates remain largely robust to this adjustment.

Finally, we conduct robustness checks by controlling for mothers' exposure to other fertility policies that could confound our measure of the One-Child Policy implementation, including the "Later, Longer, Fewer" (LLF) policies, 1.5-Child Policy, and the Two-Child Policy. First, we control for the impact of the "Later, Longer, Fewer" (LLF) policies in the 1970s. Specifically, since the individuals in our sample cover the cohorts of 1935-1966, which would also be partially affected by the LLF in the 1970s, one may deduce that the MFPC variable strongly captures the mother's exposure to the LLF. In Table A4, we conducted a robustness check by controlling for the mother's exposure to the LLF, using the policy variation of the establishment of provincial Family Planning Leading Group (e.g., Chen and

[24]The correlation coefficient of our calculated MFPC and the new MFPC is 0.99.

[25]Specifically, we estimate a probit model by regressing an indicator for the inclusion in our baseline sample on parent's characteristics (including dummies for age, gender, education level, *Hukou* status, and widowhood status), number of children, and prefecture fixed effects.

Huang, 2020; Chen and Fang, 2021). Second, in Table A6, we also control for the impact of other variation of fertility policies since the 1980s, including the 1.5-Child Policy and the Two-Child Policy. Using the enforcement variation of the two fertility policies, we add control variables to account for the impact of these policies, and the detailed variation that we have used for each policy is described in Appendix B. Taken together, as shown by both Table A4 and Table A6, the only-child effects are largely robust after accounting for these alternative fertility policies, and the alternative policies do not consistently explain parental risk behaviors in the health and finance aspects.

# 6 Heterogeneity Analysis

In this section, we further provide a set of heterogeneity analyses following the baseline findings. We first examine whether the only child effects may differ by gender of the first birth. The theoretical priors for this heterogeneity analysis could be that the only-child effects on parental risk behaviors might be more salient among the only-daughter parents, as sons are often expected to have a more significant role in supporting their elderly parents; thus the only son may ease the tendency of parental risk-avoidance compared to those parents with an only daughter. On the basis of Equation (1), we empirically test this prior by adding the indicator of whether the first birth is a boy and its interaction term with the only-child indicator. The instrumental variables for the only child variable and its interaction term with the first-boy indicator are the MFPC and its interaction term with the first-boy indicator variable.[26]

Table 8 presents the results for this heterogeneity analysis. The results reveal some evidence regarding the heterogeneity effects of having only one child by the firstborn gender. As shown by the interaction term between the only-child variable and the first-boy indicator variable, the only-child effects seem to be less salient among the only-boy parents in several behavioral dimensions, including regular drinking and the holding of commercial insurance. In other behavioral dimensions, the heterogeneity effects are not statistically significant. Therefore, the heterogeneity results are only suggestively consistent with the above conjecture.

The second heterogeneity analysis that we conduct is to examine the rural-urban differences of the only-child effects. Similarly, using the baseline IV settings, we further include the dummy variable indicating whether the parent is with urban *Hukou* and its interaction

[26] As documented by the existing literature, the gender of the firstborn can be considered as good as random (e.g., Ebenstein, 2010; Li and Wu, 2011; Lu et al., 2017; Li et al., 2022).

term with the only child indicator. Accordingly, the instrumental variables are the MFPC variable and its interaction term with the urban *Hukou* indicator. The theoretical prior for this analysis is ambiguous due to the differences in rural-urban dichotomy in various aspects. For instance, it is likely that urban parents are better supported by the pension system and health care system; hence, the underlying risk during their elderly life is more likely to be buffered by these social welfare systems compared to their rural counterparts. At the same time, however, it is also likely that other socioeconomic differences in rural and urban areas may affect the heterogeneity of the only-child effects. As such, we empirically analyze the heterogeneity effects and emphasize that the interpretation of urban-rural heterogeneity should be treated with care.

Table 9 presents the results for the urban-rural heterogeneity analysis. The results show that only-child parents in rural areas show significantly more risk-avoidance behaviors than only-child parents in urban areas. The differences are salient in both health and finance-related domains. Rural parents tend to smoke and drink less, are more likely to have fitness expenditures, and have higher saving rates compared to urban parents as a result of being the only-child parents. Moreover, in the interest of understanding the determinants of risk behaviors, having an urban *Hukou* is associated with a greater tendency in risk-avoidance behaviors in certain aspects, by being less likely to smoke and drink regularly and having a higher saving rate than the rural residents.

# 7 Mechanisms

In this section, we examine the channels through which having an only child would affect parental risk behaviors. Our analysis includes two components. In the first component, we study the only-child effects using both cultural and institutional variations regarding the importance of children in elderly support. In the second set of analyses, we further examine whether having one child can directly affect parental risk preferences.

## 7.1 Evidence from the role of elderly support

In this section, we first study the only-child effects using both cultural and institutional variations regarding the importance of children in elderly support. As suggested by existing policy discussions and anecdotal evidence (e.g., Wang, 2004; Mu, 2009), only-child parents face greater vulnerability in elderly support compared to multiple-child parents as they have only one child to rely on. In this section, we examine this mechanism by conducting two sets of analyses, both of which use the variation regarding the importance of children in elderly support.

First, we conduct a heterogeneity analysis using the regional variation in the strength of tradition regarding children's importance for elderly support. Specifically, we use the information from CHARLS 2011 with the following question: "*What can you rely on for old-age support?*" This question is answered at the individual level, and the options include: "*Children; savings; pension or retirement salary; commercial pension; and other sources.*" In our sample, 70.17% individuals indicate children as the source of elderly support, showing that the Chinese elderly anticipate that their children are an important source of support. To acquire regional variation of the strengths of the children-for-support culture, we calculate the share of individuals who indicated that children would serve as elderly support at the prefecture level. We then standardized this variable (by subtracting the mean and then dividing the standard deviation of the variable among all prefectures in the sample) so that a greater value indicates stronger preferences considering children as the main elderly support. In the empirical specification, we include an interaction term of the only child indicator and the regional importance of children in parental elderly support.[27] The instrumental variables are accordingly the MFPC variable and its interaction term with the regional intensity variable. As shown by Table 10, the only-child effects on parental risk-related behaviors vary significantly by regional preferences considering children as the main source of support for the elderly. The effects are consistently larger in regions with stronger preferences, implying that the only-child parents in prefectures with stronger traditions considering children as the main source of support for the elderly show more risk-avoidance behaviors than those with weaker traditions. The results are consistent with the channel we have speculated.

Next, we conduct another set of analyses, using the regional variation in institutional elderly support. In particular, we capture the regional variation in institutional elderly support using the size of the civil servant system by prefecture, where the system is associated with better institutional elderly support in pension and medical care schemes. To measure the size of the civil servant system across regions, we use the CHARLS data to calculate the proportion of individuals who ever worked as civil servants at the prefecture level. We standardize this variable so that a higher value indicates greater coverage of the civil servant system in the prefecture. In the analysis, we include an interaction term of the only child indicator and the local size of the civil servant system. The instrumental variables are accordingly the MFPC variable and its interaction term with the local civil servant system

[27] It should be noted that, as the variable is cross-sectional at the prefecture level, the main effect of the regional importance of children in parental elderly support is absorbed by the prefecture fixed effects.

size. As shown by the results in Table 11, we find the only-child effects on parental risk-related behaviors vary by the regional scale of the civil servant system. Specifically, the only-child risk avoidance effects are weaker in regions with a larger scale of the civil servant system. Thus, the results suggest that greater coverage of the civil servant system, which indicates better institutional elderly support, can buffer the only-child effects on parental risk behaviors.

Taken together, the above two sets of evidence show that the risk-avoidance effects of having only child are intensified in regions with stronger tradition that considers children as the main elderly support and are buffered in regions with a larger scale of civil servant system, which indicates better institutional support for the elderly. Thus, these results are consistent with the speculated channel, which underscores the vulnerability of only-child parents in terms of their elderly support.

## 7.2 Evidence from parental risk preferences

In this section, we move on to examine whether having one child would affect parental risk preferences.

Our data on risk preferences come from the China Family Panel Studies (CFPS), a nationally representative survey that covers 25 provinces in China. It is conducted by the Institute of Social Science Surveys (ISSS) of Peking University and is considered one of the most comprehensive surveys of its kind in China. It is designed to examine social and economic changes at the individual, family, and community levels. The sampling method used is PPS (probability proportional to size) and produces a sample that is representative of 95% of the population of China. The first national baseline took place in 2010, with follow-up waves every two years. CFPS 2018 elicited respondents' risk preferences using a hypothetical lottery question.

The hypothetical lottery measure of individual risk preferences in CFPS 2018 has been used in the literature (e.g., Hanaoka et al., 2018; Guiso and Paiella, 2004; Falk et al., 2018). In particular, the measure is based on five sequential binary choices, where each choice is between a fixed lottery (there is a 50:50 chance of winning 200 RMB or nothing) and varying sure payments (varying from 50 RMB to 150 RMB). In terms of the sequence, the binary choice questions take a "staircase" procedure that gradually approximates the individual certainty equivalent.[28] The individual reservation price would fall into the interval between two prices, where the first price is the sure payment where the respondent switches from

[28]The specific sequence of the questions is provided in Figure A2, and the specific questions are included in Appendix C.

choosing the lottery versus the sure payment, and the second price is the sure payment immediately before the switch. Following Hanaoka et al. (2018) and Cramer et al. (2002), we define the reservation price as the midpoint of the two prices, which provides an implicit point estimate of the measure of an individual's risk preference; for the upper- and lower-bound cases in the staircase procedure, we define the reservation price as the closest value of the sure payment.

We then use three approaches to transform the reservation price ($\lambda$) to the risk preferences. The first two measures follow Hanaoka et al. (2018) and Cramer et al. (2002). Specifically, the first measure is a simple transformation of the reservation price as follows, which we label as *Risk aversion measure* 1:

$$Transformed\ Price = 1 - \lambda/\alpha Z \tag{5}$$

where Z is the price of the lottery (200 RMB) and $\alpha$ is the probability of winning the prize (=0.5). Thus, the expected value $\alpha Z$ is 100 RMB. The values of transformed price take values between -0.5 and 0.5, and the higher value indicates greater risk aversion.

The second measure that we construct captures absolute risk aversion using the Arrow-Pratt measure, which we labeled *Risk aversion measure* 2:

$$Abosolute\ Risk\ Aversion = (\alpha Z - \lambda)/[(1/2)(\alpha Z^2 - 2\alpha Z\lambda + \lambda^2)] \tag{6}$$

In our setting, the value of absolute risk aversion ranges from -8 to 8. The higher the value of absolute risk aversion of the respondent, the more risk averse the respondent.

The third measure that we use to capture risk preference is a binary variable indicating whether the respondent is risk averse, namely, when the reservation price is less than 100. Thus, less variation is being used with this measure. In our sample, 82.7% of the respondents are risk-averse. This share is close to 78% in the German population as in Dohmen et al. (2011), and 81% in the United States as in Holt and Laury (2002).

In the analysis, our sample includes elderly parents aged 45-76 in 2018, who have married only once, with at least one child, and are not an ethnic minority. Table 12 shows the effects of having only one child on parental risk preference using the IV estimation.[29] The control variables closely follow our main specification, including gender, educational attainment, *Hukou* status, widowhood status, spousal age gap, smoking and drinking behavior before birth of the first child, region-cohort fixed effects, prefecture fixed effects, and prefecture-

[29] We also collected the FPC establishment year for the prefectures that are covered in the CFPS.

specific cohort trends. We do not control for the spousal educational gap in this specification, as our sample included the elderly group with a large portion of missing values in this variable.

Columns (1), (2), and (3) adopt the three risk preference measures, respectively. The results show that only-child parents are more risk averse than multiple-child parents, as shown in Columns (1) and (2). When using the binary measure, the effects are not statistically significant. In terms of magnitude, taking Column (1) as an example, we find that the impact of having only one child is 1.24 times its sample mean (=0.387/0.311). Together, the results show that having only one child weakly affects the risk preferences of parents. Taking into account the results in the previous sections on risk behaviors, the findings show that only child parents not only show a greater risk avoidance in health and finance behaviors, but also show stronger risk aversion in terms of their preferences.

# 8 Conclusions

This study examines the impact of having only one child on parental risk-related behaviors by exploiting the plausibly exogenous variation generated by the One-Child Policy in China. The main results show that parents who have more precarious old-age support with only one child engage in more risk avoidance behaviors, consistently in both health- and finance-related domains. Specifically, in the health-related domain, the only-child parents are less likely to smoke and drink regularly, spend more time on regular exercises, and are more likely to have fitness expenditures. Financially, only-child parents are also more likely to have commercial insurance and have a higher saving rate.

We also provide some insights into the channel through which having an only child would affect parental risk behaviors, which points to children's roles in supporting the elderly. In particular, we find that the only-child effects are intensified in regions with a stronger tradition considering children as the main source of support for the elderly, and are buffered in regions with a larger scale of civil servant system that indicates better institutional elderly support. Furthermore, we also examine whether having only one child directly affects parental risk preferences. Using data from China Family Panel Studies (CFPS) with measures of risk preferences, we find that only-child parents show stronger risk aversion in terms of their preferences. Taken together, the results indicate that only-child parents not only show greater risk avoidance in terms of risk-related behaviors, but also show more risk aversion in terms of risk preferences.

To this point, our paper is relevant to policy discussions on population aging in China, by highlighting the impact of family structure on individual risk behaviors. At the same time, we also would like to denote the limitations of our study, that the analysis cannot disentangle the role of risk perception, which could be another determinant of parental risk behaviors. We leave the exploration for future research in this direction.

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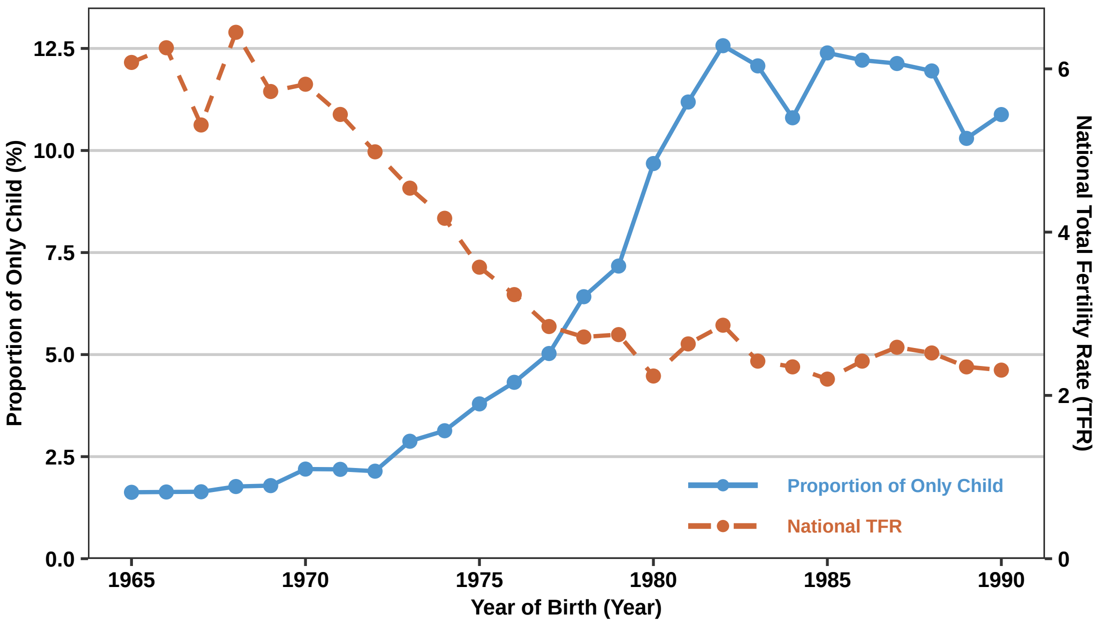
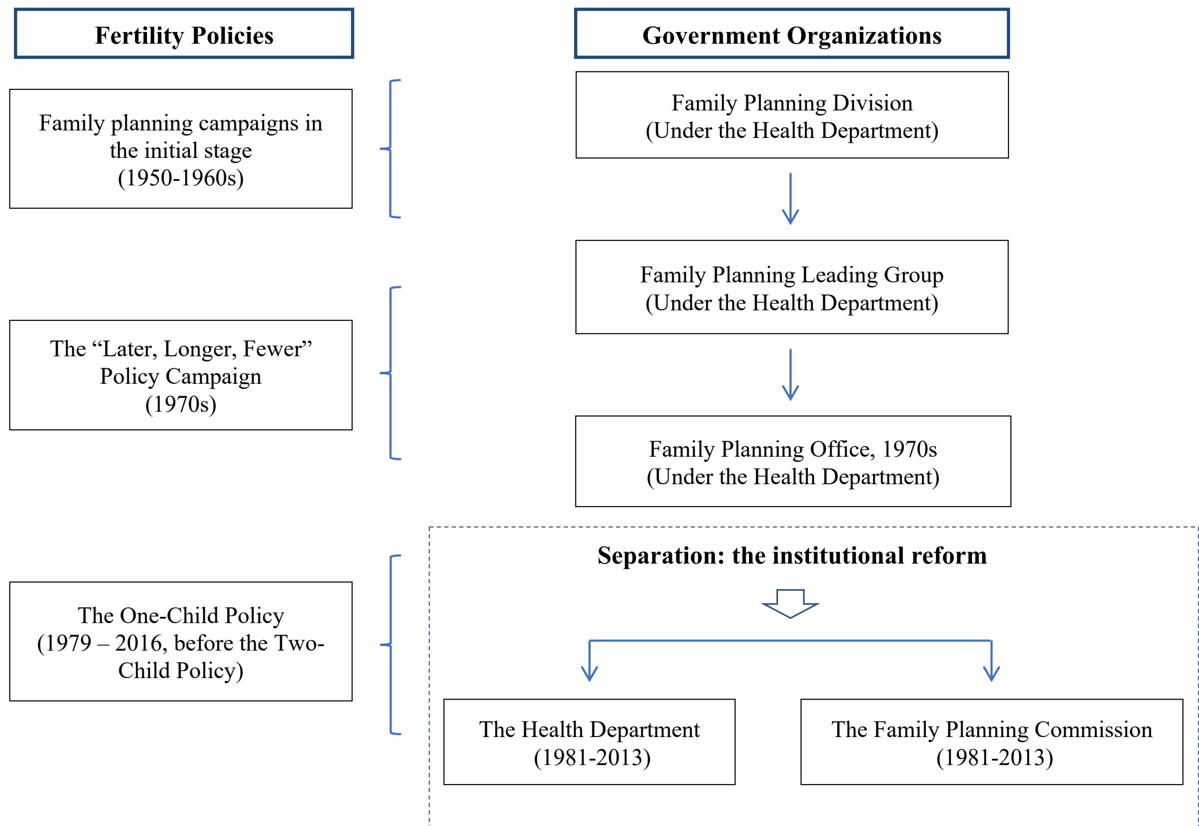
# 10 Figures and Tables

Figure 1: Trends of Only-child Proportion and National Total Fertility Rate (TFR)



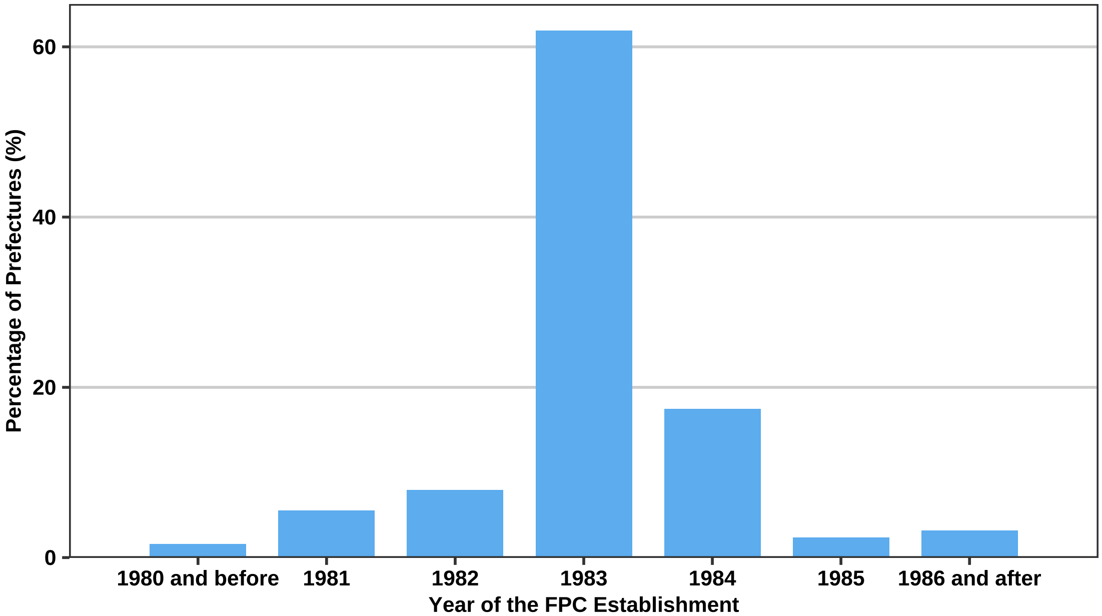

*Source*: China Population Census 2015 and Yang (2004).
*Notes*: The red dash line shows the national TFR over years (right $y$-axis) from Yang (2004). The blue solid line shows the share of only children by year of birth in the 2015 census (left $y$-axis). The share of only children is measured here from the children's perspective at the individual level. The share is calculated based on individual-level questions for married women aged 15-50, on whether the woman and/or her spouse is an only child. In the 2015 census, married women aged 15-50 were asked whether they or their spouses were only children or not. We then transform this woman-spouse-level information to individual-level information of only-child status, and the unit of observation is the woman/her spouse. Thus, the share of only-children is calculated with the numerator as the number of only children (either the woman or her spouse), and the denominator as two times the number of women who answered this question.

Figure 2: Fertility Policies and Organizational Reforms



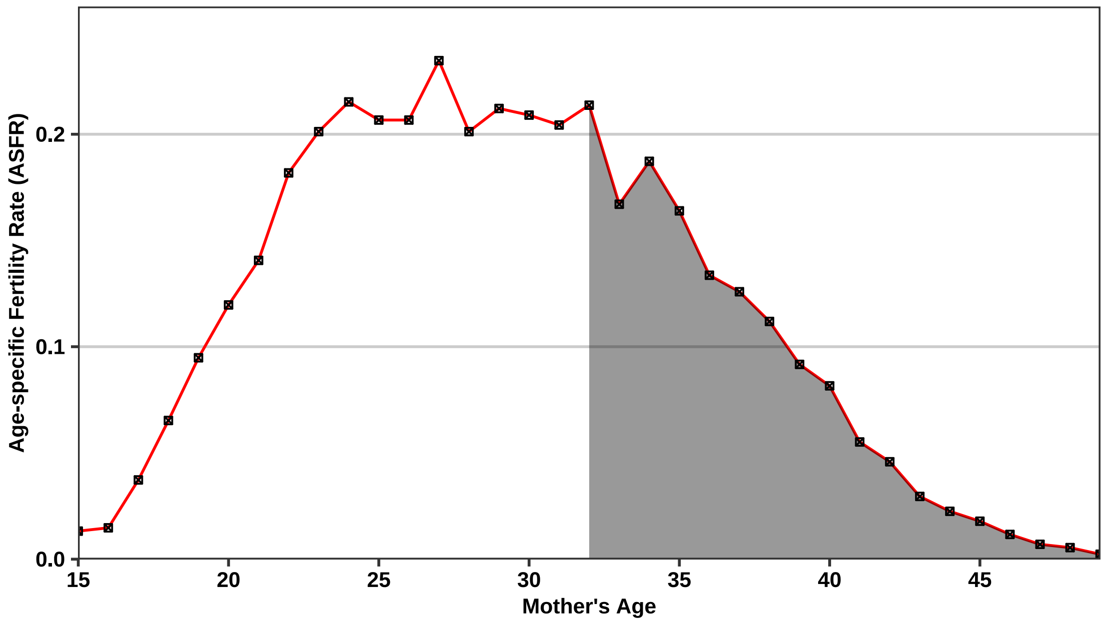

*Source*: Huang (2014); Chen (2015); Zhang (2015).
*Notes*: This figure summarizes the institutional settings of the family planning institutions in contemporary China.

Figure 3: Histogram of Prefecture-level Year of the FPC Establishment



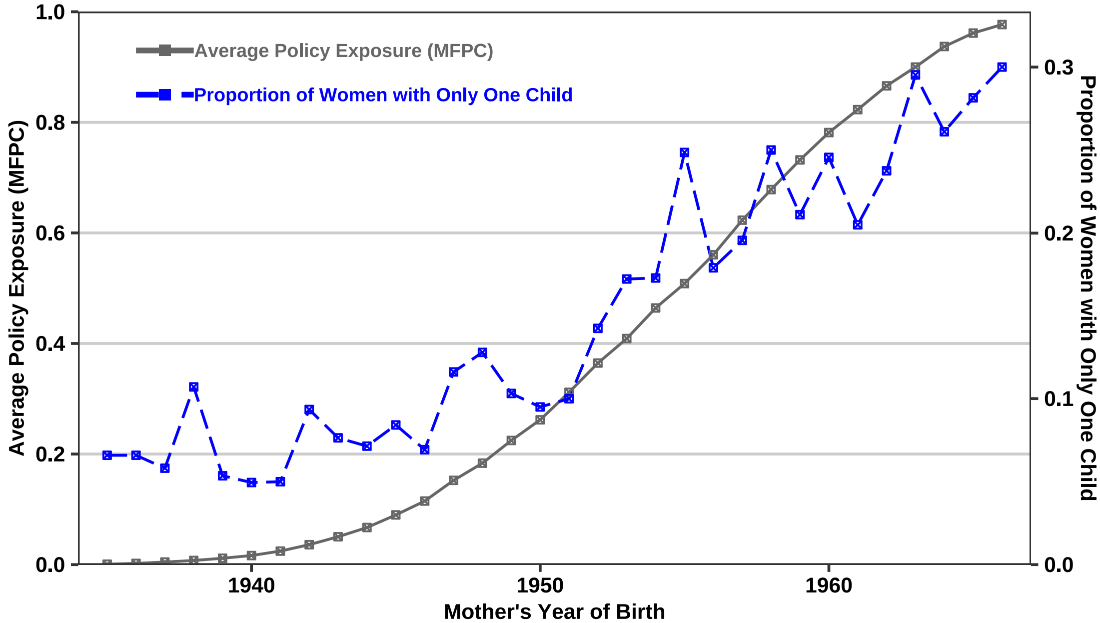

*Source*: Prefecture year books, gazetteers, and online encyclopedia including the PKU-Law dataset, and Baidu; Peng (1997).
*Notes*: The figure shows the distribution of the years of the FPC establishment in sampled prefectures (N = 126).

Figure 4: Age-specific Fertility Rate and Construction of Policy Exposure Variable MFPC




*Source*: Own construction based on CHARLS 2011.
*Notes*: The figure shows the age-specific fertility rate (ASFR) we use to construct the policy exposure variable, which is calculated using the fertility history of women born between 1925 and 1940 from the CHARLS 2011 sample.

Figure 5: Average Policy Exposure (MFPC) and Proportion of Only-Child Family




*Source*: Own construction based on CHARLS 2011.
*Notes*: The solid line indicates mothers' average policy exposure (MFPC) by birth cohort (left $y$-axis), and the dashed lined shows the proportion of only-child family by mother's birth cohort correspondingly (right $y$-axis).

Table 1: Descriptive statistics

| | N | Mean | S.D. | Min | Max |
|---|---|---|---|---|---|
| ***Panel A. Basic Characteristics*** | | | | | |
| Having only one child | 13,068 | 0.180 | 0.384 | 0 | 1 |
| Male | 13,068 | 0.489 | 0.500 | 0 | 1 |
| Age | 13,068 | 57.867 | 8.204 | 45 | 76 |
| Illiterate or semi-literate | 13,068 | 0.421 | 0.494 | 0 | 1 |
| Primary school | 13,068 | 0.224 | 0.417 | 0 | 1 |
| Middle school | 13,068 | 0.224 | 0.417 | 0 | 1 |
| High school or above | 13,068 | 0.131 | 0.337 | 0 | 1 |
| Urban *Hukou* | 13,068 | 0.209 | 0.407 | 0 | 1 |
| Widowed | 13,068 | 0.080 | 0.271 | 0 | 1 |
| Couple's age gap (husband - wife) | 13,068 | 2.164 | 3.466 | -31 | 26 |
| Husband's education is better than wife's education | 13,068 | 0.252 | 0.434 | 0 | 1 |
| Husband's education is lower than wife's education | 13,068 | 0.060 | 0.237 | 0 | 1 |
| Gender of the first child is male | 13,068 | 0.531 | 0.499 | 0 | 1 |
| ***Panel B. Risk-Related Behavioral Characteristics*** | | | | | |
| Regular smoking (currently) | 13,068 | 0.325 | 0.469 | 0 | 1 |
| Smoking history (before the first childbirth) | 13,068 | 0.262 | 0.440 | 0 | 1 |
| Regular drinking (currently) | 13,068 | 0.259 | 0.438 | 0 | 1 |
| Drinking history (before the first childbirth) | 13,068 | 0.185 | 0.388 | 0 | 1 |
| Number of days with exercise above 10 minutes in a week | 5,229 | 3.620 | 3.291 | 0 | 7 |
| Any fitness expenditure in the last year | 12,943 | 0.244 | 0.430 | 0 | 1 |
| Holding any commercial insurance | 12,966 | 0.050 | 0.219 | 0 | 1 |
| Saving rate 1 = 1 - expenditure/income | 11,209 | 0.502 | 0.701 | -3 | 1 |
| Saving rate 2 = Ln (income/expenditure) | 12,058 | 1.240 | 2.090 | -7.662 | 14.509 |
| Risk index 1 (including the exercise habit measure) | 4,363 | 0.034 | 0.411 | -1.583 | 1.434 |
| Risk index 2 (excluding the exercise habit measure) | 10,914 | 0.033 | 0.453 | -1.697 | 1.825 |
| Instrumental variable (MFPC) | 13,068 | 0.504 | 0.333 | 0 | 1 |

*Notes*: This table shows the descriptive statistics of the main variables of the parent-level data in the CHARLS 2011.

Table 2: Regional Socioeconomic Characteristics in 1977 and FPC Establishment Years

| | Dependent Variable | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Year of establishment | | | | Order of establishment | | | |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
| Share of junior high school education | 0.558 | 4.812 | 6.890 | -0.190 | 0.542 | 3.587 | 5.612 | -0.343 |
| | (3.650) | (3.960) | (4.481) | (2.900) | (3.116) | (3.311) | (3.914) | (2.769) |
| Share of senior high school education and above | -5.888 | -5.359 | -5.029 | -0.866 | -5.484 | -4.236 | -3.937 | -0.748 |
| | (4.362) | (4.514) | (5.881) | (4.981) | (4.292) | (4.005) | (5.355) | (4.766) |
| Share of secondary workers | -0.999 | -0.978 | -1.088 | 0.561 | -0.787 | -0.720 | -0.883 | 0.528 |
| | (0.915) | (0.972) | (1.361) | (0.846) | (0.864) | (0.877) | (1.240) | (0.790) |
| Share of tertiary workers | 1.583 | 0.361 | -0.660 | -0.631 | 1.404 | 0.317 | -0.563 | -0.298 |
| | (2.562) | (2.568) | (4.682) | (3.218) | (2.455) | (2.464) | (4.151) | (2.946) |
| Unemployment rate (%) | 0.336 | -1.280 | -2.252 | -0.429 | 0.511 | -0.877 | -1.784 | -0.316 |
| | (1.610) | (1.957) | (2.446) | (1.546) | (1.562) | (1.837) | (2.314) | (1.299) |
| Completed fertility rate | -0.294 | -0.225 | -0.202 | -0.263 | -0.275 | -0.206 | -0.195 | -0.204 |
| | (0.203) | (0.202) | (0.266) | (0.229) | (0.179) | (0.175) | (0.235) | (0.202) |
| GDP per capita | | -0.001 | -0.001 | | | -0.001* | -0.001 | |
| | | (0.001) | (0.001) | | | (0.001) | (0.001) | |
| Crude birth rate | | 0.037 | 0.112 | | | 0.020 | 0.078 | |
| | | (0.054) | (0.079) | | | (0.036) | (0.050) | |
| Urbanization rate | | | 0.638 | | | | 0.420 | |
| | | | (3.771) | | | | (3.146) | |
| Province fixed effects | No | No | No | Yes | No | No | No | Yes |
| Observations | 126 | 119 | 82 | 126 | 126 | 119 | 82 | 126 |
| R-square | 0.031 | 0.091 | 0.160 | 0.521 | 0.036 | 0.102 | 0.170 | 0.528 |

*Notes*: Data are from the 1982 Population Census and National Bureau of Statistics (2010). The prefecture-level socioeconomic variables include: the share of people with junior high school education among the population aged 25-50 in 1977, the share of people with senior high school education and above among the population aged 25-50 in 1977, the share of secondary workers among the population of 25-50 in 1977, the share of tertiary workers among the population of 25-50 in 1977, the unemployment rate (%) of the population aged 25-50, and the completed fertility rate of women aged 45-54 in 1977. The province-level socioeconomic variables include: GDP per capita in 1977, the crude birth rate in 1977, and the ratio of urban population to total population in 1977. Robust standard errors in parentheses are clustered at provincial level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 3: Effects of Having Only One Child on Risk Behaviors (OLS Estimates)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| ***Panel A. Control for individual-level characteristics*** | | | | | | | | | |
| Only Child | 0.005 | 0.019** | 0.288* | 0.043* | 0.032*** | 0.035* | 0.102 | 0.049*** | 0.037** |
| | (0.009) | (0.009) | (0.147) | (0.022) | (0.010) | (0.021) | (0.066) | (0.018) | (0.015) |
| Observations | 13,068 | 13,068 | 5,229 | 12,943 | 12,966 | 11,209 | 12,058 | 4,363 | 10,914 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| ***Panel B. Additional control variables*** | | | | | | | | | |
| Only Child | -0.007 | -0.002 | 0.125 | 0.011 | 0.015 | 0.004 | -0.011 | 0.010 | 0.009 |
| | (0.010) | (0.009) | (0.151) | (0.019) | (0.009) | (0.025) | (0.079) | (0.020) | (0.016) |
| Observations | 13,068 | 13,068 | 5,229 | 12,943 | 12,966 | 11,209 | 12,058 | 4,363 | 10,914 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sample mean | 0.325 | 0.259 | 3.620 | 0.244 | 0.050 | 0.502 | 1.240 | 0.035 | 0.033 |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Panels A and B report the OLS estimates when only individual-level covariates and a full set of control variables (both individual-level covariates and fixed effects) are controlled for respectively. The fixed effects controls include region-cohort FEs, prefecture FEs, and prefecture linear trends, which are added to be consistent with the IV specification for comparison. The individual-level control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 4: Effects of Having Only One Child on Risk Behaviors (IV and Reduced-form Estimates)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| ***Panel A. TSLS, 1st stage*** | | | | | | | | | |
| MFPC | 0.204*** | 0.204*** | 0.216*** | 0.202*** | 0.201*** | 0.224*** | 0.204*** | 0.238*** | 0.225*** |
| | (0.040) | (0.040) | (0.060) | (0.040) | (0.040) | (0.044) | (0.042) | (0.069) | (0.045) |
| KP $F$-statistics | 26.40 | 26.40 | 12.95 | 24.76 | 25.48 | 25.58 | 24.14 | 11.84 | 25.15 |
| ***Panel B. TSLS, 2nd stage*** | | | | | | | | | |
| Only Child | -0.455* | -0.807*** | 6.475** | 0.782*** | 0.203** | 0.891*** | 2.604** | 1.562*** | 1.266*** |
| | (0.264) | (0.250) | (2.949) | (0.247) | (0.087) | (0.320) | (1.100) | (0.532) | (0.315) |
| ***Panel C. Reduced-form*** | | | | | | | | | |
| MFPC | -0.093* | -0.165*** | 1.399*** | 0.158*** | 0.041** | 0.200*** | 0.532*** | 0.372*** | 0.285*** |
| | (0.049) | (0.039) | (0.530) | (0.036) | (0.017) | (0.063) | (0.186) | (0.076) | (0.043) |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Sample mean | 0.325 | 0.259 | 3.620 | 0.244 | 0.050 | 0.502 | 1.240 | 0.035 | 0.033 |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Panels A and B report the first and second stage results of TSLS estimation respectively. Panel C reports the reduced-form result of the effects of mother's policy exposure to the FPC (MFPC) on risk behaviors. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table 5: Falsification tests: Controlling for Experiences of Induced Abortion (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.359 | -0.719*** | 7.860* | 0.751*** | 0.182* | 0.858*** | 1.825* | 1.730** | 1.060*** |
| | (0.257) | (0.257) | (4.029) | (0.263) | (0.099) | (0.319) | (1.028) | (0.663) | (0.290) |
| Induced abortion | 0.038* | 0.056* | -0.767 | -0.095*** | -0.017 | -0.034 | 0.018 | -0.143 | -0.085** |
| | (0.022) | (0.029) | (0.495) | (0.034) | (0.013) | (0.044) | (0.137) | (0.087) | (0.038) |
| KP $F$-statistic | 28.14 | 28.14 | 9.414 | 26.48 | 27.78 | 28.69 | 25.76 | 9.840 | 27.91 |
| Observations | 10,853 | 10,853 | 4,367 | 10,760 | 10,769 | 9,338 | 10,078 | 3,655 | 9,114 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The table examines the effects of having only one child on risk behavioral outcomes when the induced abortion experiences of mothers are controlled for. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table 6: Placebo Tests: Effects of Mother's Policy Exposure to the FPC (MFPC) on Risk Behaviors (Reduced-form Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| ***Panel A. Ethnic minority individuals*** | | | | | | | | | |
| MFPC | -0.272 | 0.161 | 3.405 | -0.158 | -0.012 | -0.005 | 0.498 | -0.449 | 0.132 |
| | (0.243) | (0.241) | (2.415) | (0.169) | (0.056) | (0.330) | (0.670) | (0.357) | (0.224) |
| Observations | 736 | 736 | 247 | 734 | 729 | 601 | 653 | 204 | 594 |
| ***Panel B. Ethnic minority families*** | | | | | | | | | |
| MFPC | -0.412** | -0.071 | 1.488 | 0.012 | -0.013 | -0.053 | 0.437 | -0.313 | 0.224 |
| | (0.201) | (0.199) | (1.842) | (0.138) | (0.044) | (0.271) | (0.674) | (0.289) | (0.186) |
| Observations | 877 | 877 | 303 | 874 | 868 | 734 | 800 | 254 | 725 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to ethnic minority parents aged 45-76 who have married only once, and with at least one child. The table reports reduced-form results of the effects of mother's exposure to FPC (MFPC) on parental risk behaviors. Panel A restricts the sample to individuals who are an ethnic minority (excluding *Zhuang* ethnic group), whereas Panel B restricts the sample to individuals who themselves or their spouses are an ethnic minority. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table 7: Robustness Checks: Effects of Having Only One Child on Risk Behaviors (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking (Yes = 1) | Drinking (Yes = 1) | Days of exer. (# of days) | Fitness exp. (Yes = 1) | Com. insurance (Yes = 1) | Saving rate 1 (1-exp/inc) | Saving rate 2 Ln (inc/exp) | Index 1 (inc. exer.) | Index 2 (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| ***Panel A. Add Province-Cohort fixed effects*** | | | | | | | | | |
| Only Child | -0.439 | -0.824*** | 8.873** | 0.698*** | 0.297*** | 0.889** | 2.739** | 1.533*** | 1.401*** |
| | (0.301) | (0.255) | (3.775) | (0.252) | (0.102) | (0.374) | (1.191) | (0.568) | (0.375) |
| KP $F$-statistic | 25.36 | 25.36 | 9.075 | 23.85 | 24.65 | 21.70 | 20.74 | 9.043 | 20.04 |
| Observations | 13,014 | 13,014 | 5,128 | 12,889 | 12,911 | 11,153 | 12,004 | 4,251 | 10,855 |
| ***Panel B. Alternative source of ASFR*** | | | | | | | | | |
| Only Child | -0.433 | -0.657** | 6.026* | 0.983*** | 0.199 | 1.197** | 3.775** | 1.777** | 1.404*** |
| | (0.342) | (0.283) | (3.529) | (0.375) | (0.132) | (0.485) | (1.569) | (0.803) | (0.437) |
| KP $F$-statistic | 14.78 | 14.78 | 6.935 | 13.83 | 14.14 | 15.44 | 14.83 | 6.157 | 15.55 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| ***Panel C. Adjustment for sample selection*** | | | | | | | | | |
| Only Child | -0.200 | -0.658*** | 6.640** | 0.741*** | 0.118 | 0.540** | 2.296** | 1.759** | 0.975*** |
| | (0.221) | (0.214) | (3.059) | (0.224) | (0.087) | (0.265) | (0.975) | (0.703) | (0.251) |
| KP $F$-statistic | 27.52 | 27.52 | 10.55 | 25.44 | 26.35 | 26.73 | 25.36 | 7.826 | 25.49 |
| Observations | 12,994 | 12,994 | 5,200 | 12,870 | 12,893 | 11,143 | 11,989 | 4,335 | 10,851 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Panel A reports the results controlling for province-cohort fixed effects in the regressions. Panel B reports the results when MFPC is calculated using ASFR data in 1977 from Coale and Chen (1987). Panel C reports the results when sampling weights are applied (the weights are truncated at the 99.5th percentile to ensure that no observation is unduly weighted). Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table 8: Heterogeneous Effects by Gender of the First Birth (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.561 | -1.215** | 8.674** | 1.043** | 0.377** | 1.254** | 3.809* | 2.167** | 1.888*** |
| | (0.431) | (0.467) | (4.154) | (0.445) | (0.149) | (0.583) | (2.021) | (0.922) | (0.642) |
| First Boy | 0.026 | 0.009 | 0.188 | -0.014 | 0.008 | 0.013 | 0.149 | 0.046 | 0.020 |
| | (0.019) | (0.022) | (0.290) | (0.027) | (0.012) | (0.053) | (0.133) | (0.069) | (0.042) |
| First Boy × Only Child | 0.119 | 0.515* | -4.294 | -0.320 | -0.221** | -0.465 | -1.550 | -1.016 | -0.788* |
| | (0.234) | (0.281) | (2.675) | (0.278) | (0.109) | (0.396) | (1.284) | (0.664) | (0.429) |
| CD $F$-statistic | 9.908 | 9.908 | 6.108 | 9.427 | 9.226 | 9.763 | 8.451 | 4.899 | 9.302 |
| $p$-values of AR test of the interaction term | 0.522 | 0.003 | 0.056 | 0.158 | 0.042 | 0.195 | 0.171 | 0.005 | 0.001 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The variable *First Boy* is a dummy and takes the value of 1 for those of whom first birth is a boy, and 0 otherwise. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "CD $F$-statistic" denotes the Cragg-Donald Wald $F$-statistic on testing weak instruments. The table also reports the $p$-values of the Anderson-Rubin tests on the coefficient of the interaction term equals 0, which are robust to weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 9: Heterogeneous Effects by *Hukou* Status (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking (Yes = 1) | Drinking (Yes = 1) | Days of exer. (# of days) | Fitness exp. (Yes = 1) | Com. insurance (Yes = 1) | Saving rate 1 (1-exp/inc) | Saving rate 2 Ln (inc/exp) | Index 1 (inc. exer.) | Index 2 (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.866* | -1.483** | 9.478* | 1.581** | 0.334** | 2.088*** | 5.976** | 2.266** | 2.501*** |
| | (0.503) | (0.568) | (4.989) | (0.623) | (0.166) | (0.780) | (2.541) | (0.990) | (0.826) |
| Urban | -0.104*** | -0.101*** | -0.508 | 0.047 | 0.031* | 0.357*** | 0.958*** | 0.103 | 0.250*** |
| | (0.030) | (0.038) | (0.345) | (0.042) | (0.017) | (0.066) | (0.198) | (0.065) | (0.058) |
| Urban $\times$ Only Child | 0.594* | 0.977** | -4.907 | -1.149** | -0.195 | -1.722*** | -4.983*** | -1.231** | -1.816*** |
| | (0.345) | (0.415) | (3.011) | (0.476) | (0.126) | (0.610) | (1.895) | (0.613) | (0.634) |
| CD $F$-statistic | 5.770 | 5.770 | 3.690 | 5.625 | 5.762 | 6.075 | 5.629 | 3.767 | 6.126 |
| $p$-values of AR test of the interaction term | 0.015 | 0.000 | 0.032 | 0.000 | 0.180 | 0.000 | 0.001 | 0.000 | 0.000 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "CD $F$-statistic" denotes the Cragg-Donald Wald $F$-statistic on testing weak instruments. The table also reports the $p$-values of the Anderson-Rubin tests on the coefficient of the interaction term equals 0, which are robust to weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 10: Heterogeneous Effects by Local Preferences for Elderly Support (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking (Yes = 1) | Drinking (Yes = 1) | Days of exer. (# of days) | Fitness exp. (Yes = 1) | Com. insurance (Yes = 1) | Saving rate 1 (1-exp/inc) | Saving rate 2 Ln (inc/exp) | Index 1 (inc. exer.) | Index 2 (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.481 | -1.159*** | 8.569* | 1.104** | 0.257** | 1.384*** | 4.126** | 2.039** | 1.696*** |
| | (0.401) | (0.395) | (4.586) | (0.433) | (0.124) | (0.507) | (1.767) | (0.998) | (0.514) |
| Traditional Elderly Support × Only Child | -0.048 | -0.644** | 3.903 | 0.579* | 0.103 | 0.890** | 2.836** | 0.862 | 0.819** |
| | (0.290) | (0.285) | (3.085) | (0.331) | (0.097) | (0.422) | (1.428) | (0.913) | (0.403) |
| CD $F$-statistic | 5.403 | 5.403 | 2.577 | 5.217 | 5.346 | 5.223 | 5.403 | 1.733 | 5.809 |
| $p$-values of AR test of the interaction term | 0.818 | 0.001 | 0.049 | 0.017 | 0.384 | 0.030 | 0.015 | 0.214 | .002 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The variable *Traditional Elderly Support* denotes the standardized share of individuals who indicated children as elderly support at prefecture level (by subtracting the mean and then dividing the standard deviation of the variable among all prefectures in the sample). Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "CD $F$-statistic" denotes the Cragg-Donald Wald $F$-statistic on testing weak instruments. The table also reports the $p$-values of the Anderson-Rubin tests on the coefficient of the interaction term equals 0, which are robust to weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 11: Heterogeneous Effects by Local Civil Servant Prevalence (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.464 | -1.007*** | 7.345** | 0.946*** | 0.216** | 1.230*** | 3.600** | 1.621*** | 1.517*** |
| | (0.327) | (0.329) | (3.513) | (0.332) | (0.103) | (0.414) | (1.493) | (0.566) | (0.412) |
| Institutional Elderly Support × Only Child | 0.023 | 0.488** | -2.978 | -0.395* | -0.034 | -0.783** | -2.355* | -0.289 | -0.626** |
| | (0.200) | (0.205) | (3.011) | (0.226) | (0.088) | (0.301) | (1.200) | (0.631) | (0.269) |
| CD $F$-statistic | 8.096 | 8.096 | 4.127 | 7.793 | 7.913 | 8.377 | 7.733 | 4.405 | 8.811 |
| $p$-values of AR test of the interaction term | 0.882 | 0.001 | 0.364 | 0.033 | 0.705 | 0.003 | 0.005 | 0.643 | 0.000 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The variable *Institutional Elderly Support* denotes the standardized proportion of individuals who had ever worked as civil servants at prefecture level (by subtracting the mean and then dividing the standard deviation of the variable among all prefectures in the sample). Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "CD $F$-statistic" denotes the Cragg-Donald Wald $F$-statistic on testing weak instruments. The table also reports the $p$-values of the Anderson-Rubin tests on the coefficient of the interaction term equals 0, which are robust to weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table 12: Effects of Having Only One Child on Parental Risk Preferences (IV Estimates, 2nd Stage Results)

| | Risk aversion measure 1 | Risk aversion measure 2 | Reservation price <100 (Yes = 1) |
|---|---|---|---|
| | (1) | (2) | (3) |
| Only Child | 0.387* | 6.269* | 0.357 |
| | (0.218) | (3.518) | (0.228) |
| KP $F$-statistic | 9.188 | 9.188 | 9.188 |
| Observations | 11,876 | 11,876 | 11,876 |
| | | | |
| Individual controls | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes |
| | | | |
| Sample mean | 0.311 | 5.022 | 0.827 |

*Notes*: Data are from CFPS 2018. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, whether smoked prior to first child birth, and whether drank prior to first child birth. The values of risk aversion measure 2 are multiplied by 1000 to rescale the unit. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

# 11 Appendix

## Appendix A: Calculation of the ASFR and CFR using the CHARLS 2011

We use the CHARLS 2011 to calculate the Age-Specific Fertility Rate (ASFR) and the Complete Fertility Rate (CFR), based on mothers' fertility history. In particular, we employ the sample that includes mothers born between 1925 and 1940, whose fertility-active ages mostly cover the pre-One-Child Policy period. Using the information from the sample, the ASFR can be calculated using the following equation:

$$ASFR_a = Nchild_a/Nmother \tag{A1}$$

where $Nchild_a$ is the total number of children that mothers gave birth to at age $a$, and $Nmother$ is the total number of mothers in the sample. This ratio presents the Age-Specific Fertility Rate for each age. Further, considering the fertility-active window is 15-49, we set the ASFR below 15 and above 49 to be zero, and calculate the CFR:

$$CFR = \sum_{a=15}^{49} ASFR_a \tag{A2}$$

The Complete Fertility Rate for women of the cohort 1925-1940 can be obtained using Equation (A2).

## Appendix B: Robustness checks: Controlling Other Fertility Policies since the 1980s

In this section, we describe the specific policy variation that we have used regarding the 1.5-Child Policy and the Two-Child Policy.

**The 1.5-Child Policy.** The 1.5-Child Policy was a major policy relaxation in the 1980s following the One-Child Policy. Key variations of the 1.5-Child Policy include (i) provincial variation, that certain provinces emphasized this policy more intensively than others; (ii) rural-urban variation, that rural households with the firstborn being a girl would have an additional birth quota; and (iii) temporal variation, that the 1.5-Child Policy was mentioned in 1984 and formally implemented nationwide in 1989 (Liang, 2014).

In a similar spirit to the MFPC variable, we construct the mother's exposure variable to the 1.5-Child Policy using the above variation, which is explicitly defined as follows:

$$1.5CP_{jc} = \begin{cases} \sum_{a=15}^{49}[ASFR(a) \times I[a+c \geq 1989]]/CFR, \; if \; j \in \{1.5 \; Provinces\} \\ 0, \; if \; j \notin \{1.5 \; Provinces\} \end{cases} \tag{A3}$$

Here, $j$ indicates the province, and $c$ represents the year of birth of the mother. The 1.5 provinces include 19 provinces as shown by Table A5. $a$ ranges from 15 to 49, indicating the fertility-active ages following the definition of the fertility-active ages of women by China's National Bureau of Statistics. $ASFR(a)$ is the age-specific fertility rate (ASFR) of women of age $a$. The ASFR in our main analysis is calculated based on CHARLS 2011, using the fertility histories of women born between 1925 and 1940, which constitutes the group whose main fertility-active windows are before 1979. We used the completed fertility rate (CFR) of the same group of women to normalize the scale from 0 to 1, which is also calculated based on CHARLS 2011. $I[a + c \geq 1989]$ is an indicator variable that captures whether a woman would be exposed to the national 1.5-Child Policy in her fertility window.

Since the 1.5-Child Policy also had variation along the rural-urban dimension and the first birth-gender dimension, we add the triple interaction terms between $1.5CP_{jc}$, a dummy variable of rural *Hukou* status, and a dummy variable indicating whether the first birth is a girl. At the same time, we also account for the three variables themselves and their two-way interaction terms.

**The Two-Child Policy.** We also control for the potential effects of another policy variation during the 1980s: the Two-Child Policy. The Two-Child Policy was implemented in a parallel manner to the One-Child Policy, yet it extended the two-child birth quotas to rural households in six provinces (see Table A5). Thus, in the regression analysis, we account for the impact of the Two-Child Policy by including the triple interaction terms between (i) the MFPC variable, which is defined by Equation (2); (ii) the Two-Child Policy indicator at the provincial level (i.e., the six provinces); and (iii) a dummy variable indicating rural *Hukou* status. Thus, the triple interaction term would capture the Two-Child Policy that rural households in the six provinces after the establishment of the FPC would be permitted with two-child birth quotas. At the same time, we also account for the above three components and their two-way interaction terms.

With the above discussion, the results of the robustness check are included in Table A6 after controlling for the two alternative fertility policies. The results show that the effects of only child remain robust and that alternative policies do not consistently explain parental risk behaviors in the health and finance aspects.

## Appendix C: Survey Items on Risk Preference

CFPS 2018 used a hypothetical lottery measure to elicit respondents' risk preferences, which consists of five binary choices between a fixed lottery and varying sure payments. The specific questions are as follows:

*N1. The following questions present comparisons between two items. Please inform us of your choices. Which one would you prefer?*

N101. Risk Experiment 1:

1. A sure payment of 100 RMB;

5. Toss a coin. If the head is observed, you will get 200 RMB; if the tail is observed, you will get nothing.

[If N101=1, then keep asking N102; otherwise, skip to N104.]

N102. Risk Experiment 2:

1. A sure payment of 80 RMB;

5. Toss a coin. If the head is observed, you will get 200 RMB; if the tail is observed, you will get nothing.

[If N102=1, then keep asking N103; otherwise, skip to BB001.] (BB001 is the next module of the questionnaire.)

N103. Risk Experiment 3:

1. A sure payment of 50 RMB;

5. Toss a coin. If the head is observed, you will get 200 RMB; if the tail is observed, you will get nothing. [Skip to BB001.]

N104. Risk Experiment 4:

1. A sure payment of 120 RMB;

5. Toss a coin. If the head is observed, you will get 200 RMB; if the tail is observed, you will get nothing. [If N104=2, then keep asking N105; otherwise, skip to BB001.]

N105. Risk Experiment 5:

1. A sure payment of 150 RMB;

5. Toss a coin. If the head is observed, you will get 200 RMB; if the tail is observed, you will get nothing.

The precise sequence of questions was depicted by the "tree" logic in Figure A2.

## Appendix D: Appendix Tables

Table A1: Effects of Having Only One Child on Time-using Outcomes (IV Estimates, 2nd Stage Results)

| | Weekly hours of taking care of grandchildren | Daily duration of sleep at night (hours) | Daily duration of midday nap (minutes) |
|---|---|---|---|
| | (1) | (2) | (3) |
| Only Child | -27.129 | -1.415 | -37.860 |
| | (33.066) | (1.640) | (39.165) |
| KP $F$-statistic | 12.95 | 13.67 | 12.82 |
| Observations | 5,221 | 5,188 | 5,211 |
| | | | |
| Individual controls | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes |
| | | | |
| Sample mean | 16.668 | 6.372 | 32.013 |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The table examines the effects of having only one child on parental time-using outcomes. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p<0.01$, ** $p<0.05$, * $p<0.10$.

Table A2: Falsification tests: Heterogeneous Effects of Local Clan Strength (Reduced-form Estimates)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| MFPC | -0.052 | -0.152*** | 0.746 | 0.173*** | 0.034 | 0.160* | 0.630*** | 0.308*** | 0.266*** |
| | (0.060) | (0.045) | (0.650) | (0.049) | (0.025) | (0.088) | (0.233) | (0.094) | (0.056) |
| Low Clan Strength × MFPC | -0.096 | -0.029 | 1.630** | -0.037 | 0.016 | 0.096 | -0.234 | 0.156 | 0.045 |
| | (0.082) | (0.057) | (0.813) | (0.073) | (0.039) | (0.122) | (0.331) | (0.116) | (0.085) |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The variable *Low Clan Strength* is a dummy and takes the value of 1 if the number of genealogies per capita is below the median at the prefecture level (indicating lower density of genealogies), and 0 otherwise. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table A3: Falsification tests: Controlling for Children's Gender and Marriage Structure (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.284 | -0.534*** | 3.774** | 0.543*** | 0.164*** | 0.612*** | 1.597** | 1.101*** | 0.878*** |
| | (0.175) | (0.152) | (1.749) | (0.147) | (0.063) | (0.215) | (0.721) | (0.288) | (0.185) |
| Unmarried Son | -0.031** | -0.049*** | 0.549*** | 0.049*** | 0.001 | 0.100*** | 0.302*** | 0.117*** | 0.101*** |
| | (0.015) | (0.015) | (0.206) | (0.019) | (0.007) | (0.027) | (0.084) | (0.036) | (0.021) |
| Unmarried Daughter | -0.051** | -0.088*** | 0.576** | 0.055** | 0.021** | 0.038 | 0.139 | 0.122** | 0.111*** |
| | (0.023) | (0.022) | (0.246) | (0.023) | (0.009) | (0.033) | (0.099) | (0.048) | (0.028) |
| KP $F$-statistic | 51.06 | 51.06 | 27.73 | 48.34 | 49.43 | 46.00 | 43.90 | 24.88 | 44.88 |
| Observations | 12,970 | 12,970 | 5,184 | 12,845 | 12,868 | 11,129 | 11,971 | 4,323 | 10,837 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The table examines the effects of having only one child on risk behavioral outcomes when the children's marriage structure is controlled for. The variables *Unmarried Son* and *Unmarried Daughter* are dummies denoting whether there are unmarried sons and unmarried daughters in the household, respectively. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table A4: Robustness Check: Controlling for Mother's Policy Exposure to the LLF (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| Only Child | -0.840** | -1.596*** | 8.488** | 0.437* | 0.283*** | 0.818** | 1.031 | 1.826*** | 1.499*** |
| | (0.324) | (0.437) | (3.801) | (0.233) | (0.102) | (0.389) | (1.277) | (0.573) | (0.386) |
| MLLF | -0.123* | -0.252*** | 0.694 | -0.109* | 0.026 | -0.025 | -0.504 | 0.098 | 0.081 |
| | (0.074) | (0.084) | (0.888) | (0.057) | (0.027) | (0.114) | (0.319) | (0.166) | (0.092) |
| KP $F$-statistic | 23.09 | 23.09 | 12.57 | 22.03 | 21.83 | 21.03 | 21.19 | 14.37 | 20.48 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. The table examines the effects of having only one child on risk behavioral outcomes when mother's exposure to the "Later, Longer, Fewer" (LLF) policies is controlled for. The variable *MLLF* indicates mother's policy exposure to the establishment of the provincial Family Planning Leading Group. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

Table A5: Fertility Policies in Rural Areas by Province

| One-Child Policy | 1.5-Child Policy | | | Two-Child Policy |
|---|---|---|---|---|
| (1) | (2) | | | (3) |
| Beijing | Hebei | Zhejiang | Hubei | Hainan |
| Tianjin | Shanxi | Anhui | Hunan | Yunnan |
| Shanghai | Inner Mongolia | Fujian | Guangdong | Tibet |
| Chongqing | Liaoning | Jiangxi | Guangxi | Qinghai |
| Jiangsu | Jilin | Shandong | Guizhou | Ningxi |
| Sichuan | Heilongjiang | Henan | Shaanxi | Xinjiang |
| | | Gansu | | |
| Total: 6 provinces | Total: 19 provinces | | | Total: 6 provinces |

*Notes*: This table summarizes three distinct types of fertility policies implemented across different provinces (Peng, 1997; Li and Lin, 2023). Provinces with the One-Child Policy allow parents to have only one child. Provinces with the 1.5-Child Policy permit parents holding rural *Hukou* to have a second child if their first child is a girl. Provinces with the Two-Child Policy universally allow parents holding rural *Hukou* to have two children.

Table A6: Robustness Check: Controlling for 1.5-Child Policy and Two-Child Policy (IV Estimates, 2nd Stage Results)

| | Parental risk behaviors | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Health-related risk behaviors | | | | Finance-related risk behaviors | | | Risk avoidance index | |
| | Smoking | Drinking | Days of exer. | Fitness exp. | Com. insurance | Saving rate 1 | Saving rate 2 | Index 1 | Index 2 |
| | (Yes = 1) | (Yes = 1) | (# of days) | (Yes = 1) | (Yes = 1) | (1-exp/inc) | Ln (inc/exp) | (inc. exer.) | (exc. exer.) |
| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
| ***Panel A. Control for 1.5-Child Policy*** | | | | | | | | | |
| Only Child | -0.629 | -1.020** | 8.993** | 0.443 | 0.343* | 1.270** | 4.194* | 1.367** | 1.593*** |
| | (0.413) | (0.472) | (4.078) | (0.314) | (0.192) | (0.508) | (2.288) | (0.589) | (0.549) |
| 1.5CP × Rural × First Girl | -0.093 | -0.221 | -0.505 | -0.029 | -0.141* | -0.074 | -0.038 | -0.164 | -0.000 |
| | (0.109) | (0.135) | (1.614) | (0.104) | (0.072) | (0.198) | (0.664) | (0.263) | (0.196) |
| KP $F$-statistic | 8.974 | 8.974 | 7.112 | 8.832 | 8.713 | 10.52 | 8.104 | 7.774 | 10.63 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| ***Panel B. Control for Two-Child Policy*** | | | | | | | | | |
| Only Child | -0.621* | -1.050*** | 6.253* | 1.145*** | 0.247** | 1.473*** | 3.947** | 1.630*** | 1.715*** |
| | (0.371) | (0.373) | (3.291) | (0.426) | (0.124) | (0.507) | (1.609) | (0.589) | (0.487) |
| MFPC × Rural × Prov_2CP | -0.037 | -0.150 | 0.143 | -0.061 | -0.016 | 0.519** | 0.829 | 0.325 | 0.337 |
| | (0.114) | (0.155) | (2.396) | (0.150) | (0.055) | (0.224) | (0.781) | (0.198) | (0.204) |
| KP $F$-statistic | 15.31 | 15.31 | 10.21 | 14.29 | 14.67 | 15.63 | 14.73 | 10.14 | 15.56 |
| Observations | 13,063 | 13,063 | 5,221 | 12,938 | 12,961 | 11,204 | 12,054 | 4,353 | 10,910 |
| Individual controls | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Region-cohort FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture FE | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |
| Prefecture trends | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes | Yes |

*Notes*: Data are from the CHARLS 2011. The sample is restricted to elderly parents aged 45-76 who have married only once, with at least one child, and are not an ethnic minority. Panel A and B report results controlling for mother's exposure to the 1.5-Child Policy and the Two-Child Policy, respectively. The variable *1.5CP* indicates mother's exposure to the 1.5-Child Policy. The variable *First Girl* is a dummy and takes the value of 1 for those of whom first birth is a girl, and 0 otherwise. The variable *Prov_2CP* is a dummy indicating the provinces that implemented Two-Child Policy. The three variables in the triple interaction terms themselves as well as their two-way interaction terms are also controlled for in the regression. Other control variables include dummies for gender, educational attainment, *Hukou* status, widowhood, spousal age gap, spousal educational gap, whether smoked prior to first child's birth, and whether drank prior to first child's birth. "KP $F$-statistic" denotes the cluster-robust Kleibergen-Paap $F$-statistic on testing weak instruments. Robust standard errors in parentheses are clustered at prefecture level. *** $p$<0.01, ** $p$<0.05, * $p$<0.10.

# Appendix E: Appendix Figures

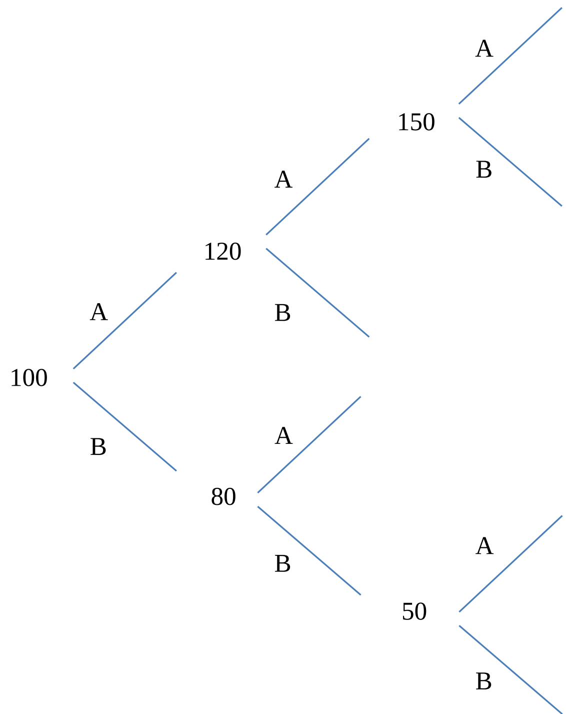
Figure A1: The History of Family Planning Institutions in Nanchang Prefecture

### 第一节　管理机构

从 1954 年起，节制生育由卫生部门专人管理。1963 年 3 月，南昌市计划生育指导委员会成立。8 月，改称南昌市计划生育委员会，下设办公室，配备有专职干部。1964 年，市属各区县先后建立相应机构。"文化大革命"中，计划生育办事机构瘫痪，1968 年 4 月，市计划生育办公室组织了临时领导小组。1971 年 9 月，恢复南昌市计划生育委员会。1973 年，根据上级要求，计划生育委员会再次调整充实。1982 年 2 月，市人民政府决定，将市计划生育委员会列为局级机构，成为市政府职能工作部门。1984 年 8 月，市政府又批准成立南昌市计划生育宣传指导所，为市计划生育委员会直属事业单位。

*Source*: Nanchang Gazetteer Compilation Committee (1997).
*Notes*: The figure above shows the history of family planning institutions and their evolution in Nanchang prefecture, Jiangxi province. As shown by the description, in February 1982, the Nanchang FPC was elevated to the bureau level within the Nanchang government, which is parallel to the health bureau. This year corresponds to the organizational shock in our study.

Figure A2: Risk Preference Tree




*Notes*: This figure shows the measurement of risk preference in the CFPS 2018. In this tree for the staircase risk task, the numbers represent a sure payment, A represents the choice of lottery, and B represents the choice of a sure payment. The staircase procedure works as follows. Initially, each respondent was asked if they would rather receive 100 RMB for certain or if they would rather take a 50:50 chance of obtaining 200 RMB or nothing. If the respondent chose the secure option ("B"), the amount of money offered in the second question was reduced to 80 RMB. On the contrary, if the respondent chose the lottery ("A"), the safe amount was increased to 120 RMB. The same logic was applied as the tree was further worked through.