The One-Child Policy in China and its Intergenerational Effects on Health*

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Abstract

We study the spillover effects of China's one-child policy on the health outcomes of subsequent generations. Despite extensive research on the effects of family size on education, few studies have examined the policy's effects on health, especially across generations. Focusing on urban Han Chinese from the China Family Panel Studies data, we use a reduced form regression discontinuity design (RDD) to isolate the local average treatment effect of the policy. The results indicate that children of policy-affected parents show significant improvements in physical and mental health, which can be attributed to increased parental investment and care and improved parental health outcomes. Our findings contribute to the literature on the intergenerational transmission of health and quantity-quality trade-offs, and highlight how family planning policies can have lasting health effects across generations.

Keywords: One-child policy, Health, Spillover Effects, Family Planning
JEL-Codes: I10, I15, J13, J18

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1 Introduction

The One-Child Policy is a family planning policy implemented nationwide in China from late 1979 to 2016 to control the country's rapid population growth. The policy strictly limited the number of children in each family to a minimum number, except those from ethnic minorities or living in rural areas. More than three decades of implementing the policy significantly changed the social dynamics and family structures in China (Settles et al., 2012; Zhang, 2017). Low fertility rates and reduced family size have led to population aging and increasing pressures on elderly care (Bai and Lei, 2020; Chen and Liu, 2009; Nie and Zhao, 2023). However, findings indicate that reducing sibling size would prompt increased investment in children, subsequently leading to improved education and health outcomes (e.g. Cáceres-Delpiano, 2006; Lee, 2008; Rosenzweig and Zhang, 2009; Zhong, 2017). Additionally, if the parental generation receives significant investment and wealth as a result of family planning policies, we expect them to allocate more resources to their offspring. Furthermore, improved health outcomes of parents can also be passed on to the next generation, resulting in better health observed among their children (Emanuel et al., 1992; Eriksson et al., 2005; Strauss and Thomas, 2007).

The existing literature focuses on the effects of sibling size and family size instead of the direct policy effects on health outcomes related to the policy's enforcement. There remains a gap regarding the intergenerational health effects of the policy itself on subsequent generations. Our paper addresses this gap by examining the intergenerational health effects of this policy, focusing on the health outcomes of individuals whose parents were affected by the policy's enforcement. By focusing on the Han Chinese population – who make up approximately 92% of China's population – in urban areas where the policy was more strictly enforced, we provide an understanding of how family planning policies can reverberate across generations. This paper also contributes to the literature by highlighting the very long-term implications of family decisions made under such policies.

We leverage data from the China Family Panel Studies (CFPS), a nationally representative biennial longitudinal survey funded by the Chinese government and conducted by the Institute of Social Science Survey (ISSS) at Peking University since 2010. The CFPS covers both economic and non-economic aspects of the Chinese population, providing rich data on economic activities, education, family dynamics, and health (Xie and Hu, 2014). We focus on three physical health measures, including the likelihood of being sick, self-rated health, interviewer-rated health and one mental health indicator – distress level, which we construct based on the Kessler Psychological Distress Scale (K6) and the Center for Epidemiologic Studies Depression Scale (CES-D)[1]. We employ a reduced form regression discontinuity design (RDD), exploiting the policy cut-off in 1980 that

[1]Detailed description of the data and our selection of outcome variables are presented in Section 3.

creates a discontinuity in the number of single-kid families, thereby precisely isolating the policy's local average treatment effect (LATE) on the next generation's health outcomes. We show that our design passes multiple checks to ensure the validity of core RDD assumptions. Our results suggest that children born to policy-affected parents, especially policy-affected mothers, demonstrate better physical and mental health. Finally, we provide empirical evidence that increased investment in children's health and improved health outcomes in the parents' generations are mechanisms driving our results. Additionally, these parents are less demanding and more responsive toward their children, which explains the lower levels of distress observed among their children.

Our baseline results, in section 5, show that children born to policy-affected mothers exhibit significant improvements in both physical and mental health. Specifically, the likelihood of them being sick decreases by 1.8 percentage points. Their self-rated health improves by 8.2 percentage points, representing an increase of 20% over the mean. Interviewers also observe better health among these children, with observed health improved by 1.8 percentage points. In addition, they experience lower levels of mental distress, with the probability of having distress reduced by roughly two-thirds. We find similar but less statistically efficient results when examining children whose fathers were born after 1980 and affected by the policy. Children born to policy-affected fathers are less likely to be sick and have better interviewer-observed health. They also have lower levels of distress, although the estimate is not statistically significant. Given the significantly smaller sample size for father data and numerous evidence showing that urban Han mothers benefited from the demographic pattern created by the one-child policy (Fong, 2002; Veeck et al., 2003; Zhang, 2019), we focus primarily on the effects of the policy from the mothers' side[2]. We conduct a wide range of sensitivity checks to show that our baseline estimates from the mothers' data remain robust.

In Section 6, we empirically investigate several mechanisms to explain our results. First, higher investment in children's health, when family size becomes smaller, can explain better health in children. This is consistent with the quantity-quality trade-off, formulated by Becker (1960), illustrating a negative correlation between family size and the resources allocated to each child. Second, the intergenerational transmission of health and household characteristics is another mechanism that elucidates our results. A lower fertility rate, which is transmitted from grandparents to parents (Kolk, 2014; Murphy and Knudsen, 2002), leads to increased human capital investment per child, supporting our narratives on child health investment. Additionally, mothers affected by the policy exhibit better health outcomes, which can pass on to their children (Emanuel et al., 1992; Eriksson et al., 2005). Third, parenting practices and parent-child interactions show that policy-affected parents demonstrate high responsiveness but low parental

[2] Although analysis using fathers' data yields similar results, given the smaller sample size and narrower bandwidth, we interpret these results with caution. See Section 5.1 for more details.

demand toward their children, resulting in their children becoming more relaxed and exhibiting lower levels of distress. This aligns with literature on parental demands (Lo et al., 2020; Soysa and Weiss, 2014; Wong et al., 2019) and parental responsiveness (Davidov and Grusec, 2006; Miller-Slough et al., 2018), especially in the Chinese context where children are the only child (Liu et al., 2010; Lu and Chang, 2013).

Related Literature. Our research question is relevant to several strands of literature. The theoretical basis for the quantity and quality of children is commonly referred to as the quantity-quality trade-off. This framework was first theorized in the work of Becker (1960), who considered children a consumption good, requiring a family to decide not only on the number of children but also on the corresponding expenditure allocated to them. This theory was further developed in the work of Becker and Lewis (1973), Becker and Tomes (1976), and Willis (1973), which emphasized the negative correlation between the quantity and quality of children due to both the "price effect" and the "income effect", given the limited resources and time a family has to invest in its children – budget and time constraints. In essence, this trade-off arises from the fact that parents have to spread their time and resources more thinly as the number of children increases (Hanushek, 1992). This model is consistent with the resource dilution model in sociology, which demonstrates that as the number of children in a family increases, resources are divided among them, resulting in each child having fewer resources and, consequently, lower quality of life (Blake, 1981, 2022).

The literature presents mixed empirical evidence on the quantity-quality trade-off. Numerous studies have found a negative association between family size and investment in children (Cáceres-Delpiano, 2006; Chen, 2020; Lee, 2008; Li et al., 2008; Ponzo and Scoppa, 2022; Rosenzweig and Zhang, 2009). However, several others have observed no evidence of the quantity-quality trade-off (Angrist et al., 2010; Black et al., 2005; Diaz and Fiel, 2021) or even a positive relationship (Gomes, 1984; Lao and Lin, 2022; Qian, 2009). The majority of these studies focus on educational attainment and schooling as a quality indicator, with only limited literature examining health outcomes as a determinant of child quality.

Several studies provide evidence of a negative association between family size and health, consistent with the quantity-quality trade-off. Liu (2014) and Zhong (2017) investigate the one-child policy in China as an exogenous shock and find negative impacts of family size on child height. A similar study conducted by Liang and Gibson (2018) considers nutrient intake as a measure of parental investment in children and discovers that an additional sibling reduces nutrient intake by between one-tenth and one-fifth of the recommended level. Chen (2021) exploits the two-kid policy in Vietnam and shows that having another sibling worsens the health of children in terms of height and weight. Moreover, using twin data, Glick et al. (2007) indicate that unplanned births in

Romania have negative effects on children's human capital, measured through nutrition and schooling and that these effects extend significantly to later-born siblings of first-born twins. Similarly, Rosenzweig and Zhang (2009) examine the effects of twinning in China and provide evidence that an additional child leads to a significant decline in self-assessed health of all children within the family.

However, some studies have found positive impacts of having siblings on children's health. Lordan and Frijters (2013) utilizing data from the Young Lives Project (YLP) in Peru find that the association between family size and health outcomes such as height is negative specifically for unplanned pregnancies, while it becomes positive for planned childbirths. Datar (2017) investigates the relationship between family size and obesity in the US and provides evidence that children with siblings have lower BMI and are less likely to be obese because they have healthier diets. Meanwhile, Millimet and Wang (2011) employing data from the Indonesia Family Life Survey observe only modest evidence. Zhong (2014) even finds null evidence of the trade-off in terms of the height and BMI using China's one-child policy, while Zhang et al. (2020) find being raised in a one-child family increases the probability of being overweight or obese.

In terms of subjective wellbeing, several studies have found being an only child negatively impacts self-reported psychological health (Wu, 2014; Zeng et al., 2020). Cameron et al. (2013), leveraging the one-child policy in China, provide evidence that these "little emperors" are less trustworthy and more pessimistic. However, Liu et al. (2010) and Rao et al. (2024) compare single-kid and multiple-kid families and show that single kids reported lower levels of psychological distress and mental health problems, as a result of higher parental responsiveness.

Our paper also relates to the literature on the intergenerational transmission of health and household characteristics across generations. Studies consistently demonstrate modest yet persistent effects in the transmission of parents' fertility patterns to their children (Kolk, 2014). Several papers have found that parents' fertility or family size preferences influence the preferences of their offspring (Anderton et al., 1987; Johnson and Stokes, 1976; Murphy and Knudsen, 2002). Another body of work focuses on investigating intergenerational transmission of health from parents to their off-spring. These studies are commonly rooted in the "fetal origin hypothesis" formulated by Barker (1990), which asserts that early-life (in-utero) health and circumstances play a crucial role in shaping health and economic conditions in later stages of life. A different framework is summarized by Strauss and Thomas (2007) who emphasize the transformation of health inputs into health outputs, given technological and biological constraints, as the mechanism explaining health transmission within a family. In particular, in addition to genetic endowments that would be transmitted across generations, non-genetic aspects of parental health also influence their ability to manage inputs into the health production function of their children. These frameworks have been validated by the work of Emanuel et al.

(1992) and Eriksson et al. (2005), which show a robust correlation between the health of parents and that of their offspring. There is also a large and growing body of literature exploring the impact of external health shocks on the health of subsequent generations (e.g. Camacho, 2008; Islam et al., 2017; Moyano, 2017).

Contribution. Our primary contribution is to provide new causal evidence about the intergenerational effects of family size and family planning policies on subsequent generations. First, we leverage data from the China Family Panel Studies (CFPS), a nationally representative and one of the most comprehensive social panel surveys conducted in China (Xie and Hu, 2014), ensuring a high level of reliability and coverage. Second, we examine the spillover effects on both the physical and mental health of children whose parents were affected by the policy. Third, we also explore several mechanisms from the CFPS data to explain our results, contributing to the literature on quantity-quality trade-offs and intergenerational effects across generations. In addition, we further address the question of parenting behaviors and parent-child relationships in modern Chinese families.

The paper is structured as follows: Section 2 provides historical background and the introduction of the one-child policy in China; Section 3 describes the data and sample; Section 4 illustrates empirical strategy and identification assumptions; Section 5 presents estimation results and robustness tests; Section 6 investigates possible mechanisms that can explain our results and Section 7 concludes the paper.

2 Background

Since 1949, China started its industrialization process, experiencing substantial population growth, with the belief that this growth would contribute to the national effort (Zhu, 2012). However, consistent poverty and high fertility rates caused fears of overpopulation (De Silva and Tenreyro, 2017). From 1970 onwards, citizens were encouraged to marry at a later age because of the large population, and in the early 1970s, the state introduced a series of birth planning policies. In 1978, the authorities began to encourage one-child families, and in early 1979 they announced their intention to advocate for one-child families, which later became a national policy.

The one-child policy was introduced in China in 1979 as a strict family planning policy to curb the country's rapid population growth (Wang et al., 2016), and it was formally written into the country's constitution in 1982. The policy focuses on the Han Chinese, which makes up 92% of the population (Huang et al., 2016b). In principle, a couple was only allowed to have one child from late 1979, except in some rural ethnic minority areas such as Xinjiang, Yunnan, Ningxia and Qinghai.

The evolution of the one-child policy had several phases (Greenhalgh, 2008; Scharping, 2013a). It was first announced as “Best is one, at most two; eliminate third births” in the second half of 1978. In December 1979, the National Population and Family Planning Commission announced the policy as “Best is one”. From February 1980, it quickly changed to “One for all” policy. Over time, however, various exceptions were made and the policy was further revised in early 1989 to “One child with exceptions for rural couples with only a daughter”.

In 2016, China officially relaxed its one-child policy, marking a significant shift in its approach to population control. The policy change, which allowed families to have two children from 2015 with some modifications, reflected growing concerns about the policy’s negative demographic and socio-economic impacts. The relaxation aimed to address issues such as the rapidly aging population, shrinking workforce, and gender imbalances. However, more urban families choose to have only one child spontaneously because of the financial and social pressures, even after the relaxation (Qian and Jin, 2024). The long-term effects of this policy change on family dynamics, economic stability and population health remain areas of policy debate.

Social consequences of the one-child policy (OCP). The one-child policy in China has led to several social consequences, notably a significant decline in the fertility rate, which had already been decreasing due to earlier family planning campaigns in the 1970s (Feng et al., 2014). The average family size reduced from 4.8 in the early 1970s to 3.1 in 2010 (Aird, 1983; Census Office of the State Council, 2020). Single-child families became prevalent, especially in urban areas, where about 80 percent of families consisted of three members by the end of the 20th century (Tu, 2016). Other direct outcomes included a skewed male-to-female ratio and higher fertility rates in rural areas compared to urban ones, disadvantaging rural families economically (Ebenstein, 2010; Hannum, 2003). Despite criticisms, urban daughters often benefited from the policy, receiving more family resources and achieving higher educational attainment and empowerment (Fong, 2002; Huang et al., 2016a, 2021). Long-term impacts include accelerated population aging, increased pressure on elderly care, and the rise of "empty nest" families in urban areas (Bai and Lei, 2020; Chen and Liu, 2009; Nie and Zhao, 2023; Yuesheng, 2014; Zhu and Walker, 2021).

Policy effects on the first generation. The first generation subjected to the one-child policy experienced notable benefits, particularly in urban areas. Families were able to concentrate their resources on their single child’s education and health, leading to substantial investments in these areas (Zhang, 2019). This focus resulted in higher educational attainment for females and improved overall health outcomes (Fong, 2002; Huang et al., 2016a; Rao et al., 2024). Women born under the one-child policy achieved higher educational levels (Huang et al., 2016a), and stricter early-life fertility restrictions increased female empowerment, as evidenced by a rise in female-headed households (Huang et al., 2021).

The policy also brought qualitative changes in family dynamics, including simplified family structures, reduced patriarchy in daughter-only families, and greater individual choice regarding family living arrangements and childbearing (Fong, 2002; Shi, 2017). Furthermore, greater parental involvement in childcare led to improved parent-child interactions (Short et al., 2001). As a result, this generation enjoys higher income levels and reduced overall stress. The persistence of intergenerational income in urban China also highlights the lasting economic impacts of the policy (Yi, 2016).

Policy cut-off in this paper. We examine the introduction of the policy during the 1979-1980 period to identify the exact policy cut-off for our study. In China, before the birth of the one-child policy, the government had imposed restrictions on the number of children a couple could have. In 1977 and 1978, both urban and rural couples were required to limit their family size to only two children (Hardee-Cleaveland and Banister, 1988). In early 1979, several intentions for the universal one-child policy were introduced, and the policy "Best is One" was officially announced in December 1979. "One for all" policy followed quickly in February 1980, clearly stating that every couple is only allowed to have one kid. Later on, the Chinese Communist Party's Central Committee issued a public letter urging all party members and the Communist Youth League to adhere to the one-child policy on 25th September 1980, a date often mentioned as the policy's "official" start date (Scharping, 2013b). The revised 1980 Marriage Law, ratified during the Third Session of the National People's Congress on September 10, 1980, also explicitly mandated that all couples must practice birth control (Hardee-Cleaveland and Banister, 1988; Hare-Mustin, 1982; Santana Cooney et al., 1991). Before that, however, strict fines for violating the one-child policy already started to be imposed nationwide in January 1980 (Santana Cooney et al., 1991). During this period, abortions were required in several provinces in China, even in the second and third trimesters of pregnancies. The number of induced abortions increased sharply in 1979 and rose even higher in subsequent years (Hardee-Cleaveland and Banister, 1988).

We expect that there was a sharp increase in the proportion of single-child births in 1980 (the first quarter of 1980 according to our data structure we will mention later). Although official announcements and legislation related to the one-child policy and birth control were issued in September 1980, the policy started to be strictly enforced in 1980 with a wide range of rigorous measures like required abortions and birth control practices. Additionally, beginning in 1980, penalties were imposed on women who had a second child without official permission (Hardee-Cleaveland and Banister, 1988), confirming that it was difficult for Han mothers living in urban areas to have another child. We later verify this cut-off date by illustrating the discontinuity in the ratio of individuals with no siblings at our predicted cut-off point (the first quarter of 1980) within our dataset.

3 Data and Descriptive Statistics

3.1 Data

We use the China Family Panel Studies (CFPS) from the Institute of Social Science Survey (ISSS) at Peking University, China[3]. CFPS is a nationally representative, biennial longitudinal survey of Chinese families and individuals, starting in 2010. For our analysis, the CFPS dataset has several key features. First, it allows us to identify the intra-household relationships, which we need for our empirical strategy using parents' birth information. Second, adults were asked about their parents' information and their siblings in the first wave – 2010, providing valuable data for studying sibling size and policy effects. Third, with six waves up to 2020, the CFPS provides national-level information on family dynamics and health outcomes for our study.

Children's health status. We use the following outcomes as measures of children's health status:

1. Whether a child was ever sick in the last month: CFPS interprets sickness as a situation in which the child experiences physical discomfort and needs to take treatment (medicines or others).
2. Children's self-rated health: The survey asked respondents to rate their health status on a scale from 1 to 5 indicates healthy, fair, relatively unhealthy, unhealthy, and very unhealthy respectively. Only those aged 10 and above answered this question. We recode this into a binary variable with 1 for healthy and 0 for all other ratings.
3. Interviewer-observed child health: The interviewer from the ISSS recorded their assessment of the health of the presented respondent, choosing from 1 (worst) to 7 (best). We also recode this variable into a binary format, with 1 for observed health rating greater than or equal to 4, otherwise 0.
4. Distress indicator based on K6 and CES-D: We incorporate two psychological scales – Kessler Psychological Distress Scale (K6) (Kessler et al., 2002) and the Center for Epidemiologic Studies Depression Scale (CES-D) (Radloff, 1977) – to measure children's mental health because these scales are available in different waves. A detailed description of the mental health scales can be found in the Appendix A. We construct a consistent variable called "distress", which is coded 1 if a child shows

[3]The data are from the China Family Panel Studies (CFPS), funded by the 985 Program of Peking University and carried out by the Institute of Social Science Survey of Peking University.

signs of mental distress, i.e. K6 is greater than or equal to 5 (Prochaska et al., 2012), CES-D8 is greater than or equal to 7, or CES-D20 is greater than or equal to 16 (Bi et al., 2023).

Mechanisms. We will later examine several potential mechanisms through which the effects of the policy on the first generation, now parents, could be passed on to the second generation, their children. First, we look at family income and expenditure, focusing on expenditure directly on the child. This may influence the family resources available for the child's well-being. Second, we examine the policy's impact on parents' health status, the number of siblings they have, their fertility choices – the number of children they decide to have, and their education level. Together, these factors shape the environment in which children grow up and can affect their upbringing. Lastly, we look at the interactions between parents and children, which can affect children's mental health. The quality and nature of these interactions are crucial in determining the emotional and psychological well-being of children. By studying these mechanisms, we aim to understand how policy effects are transmitted across generations.

3.2 Sample Selection and Summary Statistics

We compile data from the six waves of CFPS based on information from the child questionnaires to create a repeated cross-sectional dataset and merge it with data on parents and family from the corresponding questionnaires. We first drop observations missing key demographic characteristics such as children's age, gender, and rural or urban residence, birth information of both parents, and family size (around 1% dropped at this stage). We further split this dataset into two datasets where either mother's or father's information is available. For each dataset, we exclude observations from provinces with fewer than 50 mothers/fathers (5 same provinces for both). Next, we restrict the sample to Han ethnicity (87.11% of mothers, 88.01% of fathers) and urban parents (51.56% of mothers, 53.79% of fathers after keeping only Han ethnicity) due to the policy focus.

These steps ensure complete information on the birth years and months of urban Han parents. Due to the specific focus on urban Han parents, which limits the number of observations, we construct our data based on quarterly birth information for these parents to ensure statistical power. We then exclude those with missing information on children's health outcomes at each wave, resulting in different sample sizes for various outcomes. For each dataset using either mothers' or fathers' information along with their children's data, our final sample size is approximately 4,000 observations. Table 1 presents summary statistics for our sample. Most of the statistically significant differences between the control and treatment groups are in age. This is due to the nature of the policy, so that those born later, i.e. younger, are more likely to be treated.

Table 1. Descriptive statistics

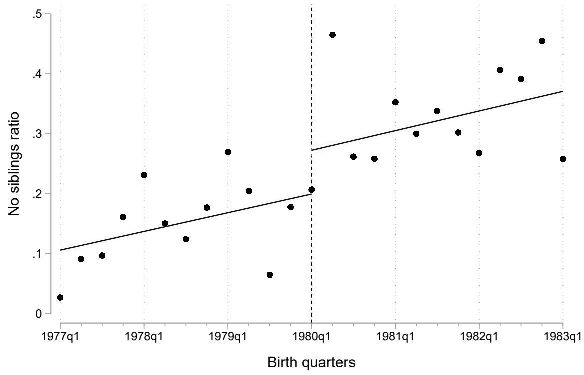
	Mothers				Fathers			
	All	Control	Treated	Diff	All	Control	Treated	Diff
A. Health Outcomes								
Child was ever sick last month	0.26	0.25	0.27	-0.02	0.28	0.28	0.28	-0.00
	(0.44)	(0.43)	(0.45)	(-1.70)	(0.45)	(0.45)	(0.45)	(-0.02)
Self-rated health (healthy = 1)	0.38	0.38	0.38	-0.01	0.39	0.37	0.41	-0.04
	(0.49)	(0.48)	(0.49)	(-0.24)	(0.49)	(0.48)	(0.49)	(-1.39)
Interviewer-observed health (≥ 4 on 1-7 scale)	0.98	0.98	0.98	-0.00	0.98	0.98	0.97	0.00
	(0.14)	(0.15)	(0.13)	(-0.56)	(0.15)	(0.15)	(0.16)	(0.47)
Distress indicator based on K6 and CESD	0.12	0.14	0.10	0.04*	0.11	0.11	0.11	0.00
	(0.33)	(0.34)	(0.30)	(2.42)	(0.31)	(0.31)	(0.31)	(0.11)
B. Demographics								
Child's age	8.29	9.23	7.48	1.76***	7.74	8.36	7.15	1.21***
	(4.26)	(4.18)	(4.16)	(13.96)	(4.19)	(4.30)	(3.99)	(9.42)
Child's gender (female = 1)	0.48	0.47	0.50	-0.03	0.49	0.47	0.50	-0.03*
	(0.50)	(0.50)	(0.50)	(-1.74)	(0.50)	(0.50)	(0.50)	(-2.23)
Child's birthyear	2006	2005	2007	-2.36***	2007	2006	2008	-1.62***
	(3.94)	(3.89)	(3.64)	(-20.68)	(3.75)	(3.89)	(3.43)	(-14.31)
Mother's age	34.75	36.16	33.51	2.65***	33.63	34.79	32.51	2.28***
	(3.64)	(3.36)	(3.42)	(25.46)	(4.32)	(4.25)	(4.09)	(16.69)
Mother's birthyear	1980	1978	1981	-3.21***	1981	1980	1983	-2.70***
	(1.84)	(0.81)	(1.00)	(-117.88)	(3.18)	(3.03)	(2.73)	(-29.49)
Father's age	36.93	38.08	35.89	2.19***	35.26	36.56	33.99	2.57***
	(4.50)	(4.48)	(4.26)	(15.13)	(3.68)	(3.47)	(3.41)	(23.27)
Father's birthyear	1978	1976	1979	-2.95***	1980	1978	1981	-3.01***
	(3.47)	(3.29)	(3.01)	(-29.77)	(1.75)	(0.80)	(0.96)	(-110.72)
Family size	4.82	4.73	4.90	-0.17**	5.17	4.84	5.48	-0.64***
	(1.75)	(1.74)	(1.76)	(-3.16)	(1.90)	(1.59)	(2.12)	(-11.13)
C. Grandparent's characteristics								
Grandfather's age	59.03	61.37	56.87	4.50***	58.61	60.71	56.46	4.25***
	(6.14)	(6.18)	(5.26)	(24.62)	(5.76)	(6.00)	(4.60)	(23.16)
Grandmother's age	60.95	62.72	59.35	3.37***	61.09	63.00	59.13	3.87***
	(6.03)	(6.16)	(5.45)	(18.07)	(6.12)	(6.00)	(5.60)	(19.23)
Literacy (grandfather)	0.83	0.78	0.87	-0.09***	0.83	0.85	0.80	0.05***
	(0.38)	(0.41)	(0.33)	(-7.46)	(0.38)	(0.36)	(0.40)	(3.98)
Literacy (grandmother)	0.63	0.57	0.69	-0.12***	0.59	0.59	0.58	0.01
	(0.48)	(0.50)	(0.46)	(-7.67)	(0.49)	(0.49)	(0.49)	(0.81)
Unemployment (grandfather)	0.14	0.12	0.16	-0.04***	0.14	0.11	0.18	-0.07***
	(0.35)	(0.32)	(0.36)	(-3.39)	(0.35)	(0.32)	(0.38)	(-5.43)
Unemployment (grandmother)	0.25	0.27	0.23	0.04**	0.24	0.23	0.25	-0.02
	(0.43)	(0.44)	(0.42)	(2.98)	(0.43)	(0.42)	(0.43)	(-1.61)
Either of grandparents is communist	0.16	0.16	0.16	-0.00	0.15	0.13	0.17	-0.04***
	(0.37)	(0.36)	(0.37)	(-0.20)	(0.36)	(0.34)	(0.38)	(-3.83)
Observations	4418	2054	2364	4418	4206	2058	2148	4206

Note: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. The table provides the mean/standard deviation of the corresponding variables within 3 years (12 quarters) around the policy cut-off date (January 1980). "All" means the whole sample, "Treated" means mothers/ fathers have no siblings, and "Control" means mothers/fathers have siblings. "Diff" shows the mean difference between treated and control groups. The scale for interviewer-observed child health ranges from 1 (worst) to 7 (best).

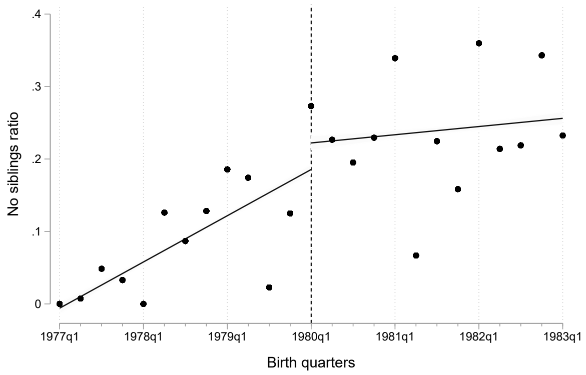
We further demonstrate the validity of the policy cut-off date in our data. Figure 1a shows the proportion of Han urban adults born within 12 quarters of the 1980Q1 policy cut-off who have no siblings. As noted above, adult sibling information is only available in the first wave – 2010. We use weighted 2010 adult data with the sampling weights provided by the CFPS to ensure national representativeness. The graph shows the simple overall national average of the no sibling ratio in each quarter. The graph presents an upward general trend in the proportion of adults without siblings immediately following the policy, and suggests a jump in this ratio in 1980Q1 among Han urban residents. Looking at females and males separately in Figures 1b and 1c, there is a jump in both figures in 1980Q1. The no sibling ratio for females shows an upward trend but starts very low before the policy was implemented and remains fairly stable afterwards, while for males the no sibling ratio is stable before but increases after the policy was implemented. As noted above, there were exceptions, such as allowing a second child when the first was a daughter, which can partly explain the different patterns for men and women.

Figure 1. Proportion of adults without siblings born within 12 quarters of policy cut-off

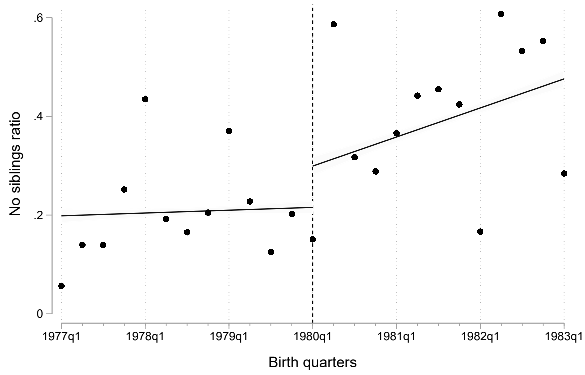
(a) Overall



(b) Female



(c) Male



Notes: Figure (a) shows the proportion of Han urban adults nationally (using sampling weights provided by the CFPS) born within 12 quarters of the policy cut-off with no siblings, while Figures (b) and (c) show the no siblings ratio for females and males respectively. The policy cut-off is 1980Q1. We use the 2010 CFPS adult data as this is the only wave that provides information on adults' siblings.

4 Empirical Strategy

4.1 Main Specifications

To explore the spillovers of the policy, we employ a regression discontinuity design (RDD) with the number of quarters between mothers' date of birth and policy date as the running variable. In an RDD, the running variable plays a crucial role in determining the treatment status when there is a discontinuity in the treatment at a specific cut-off point, such as the first quarter of 1980 in our context. However, all other covariates should exhibit smoothness at the cut-off.

In this paper, as seen in Figure 1, the no sibling ratio jump at the cutoff is actully very small, we rely on a non-parametric reduced form RDD approach to estimate the policy effects on the health outcomes of children born to parents exposed to the policy. In all regressions, we employ a triangular kernel weighting function where the weight assigned to each observation decreases as the distance from the cut-off increases. In addition, we estimate the effects from mothers and fathers separately.

For bandwidth selection, Table 2 shows the data-driven optimal bandwidths for different health outcomes following Calonico et al. (2020). We select the bandwidth equals to the average of optimal bandwidths generated for these outcomes within the parent's gender. In particular, we choose a bandwidth of 11 quarters when examining the effects from the mothers' side (maternal effects) and a bandwidth of 9 quarters when examining the effects from the fathers' side (paternal effects). All tables in our paper will show the estimates using 11-quarter bandwidth for maternal effects and 9-quarter bandwidth for paternal effects unless it has been specified differently. Specifications with other bandwidth choices will be considered in our robustness checks.

Table 2. Optimal bandwidths for RDD

	Mother's information	Father's information
	(1)	(2)
Sick	9.850	10.661
Self-rated health status	10.047	6.598
Observed health status	11.835	7.763
Distress	12.446	9.642

Notes: The table shows the mean square error optimal bandwidths of main outcomes following Calonico et al. (2020) – CCT bandwidths with a local linear polynomial. Standard errors are clustered at the mothers' or fathers' birth years. Column (1) presents the optimal bandwidth for each outcome using mothers' information while column (2) using father's information. The average optimal bandwidth using mother's information is around 11 quarters, and the average optimal bandwidth using father's information is around 9 quarters.

The regression measuring the direct effect of the policy on the first-generation or parents' outcomes takes the following form:

$$\begin{aligned} Y_i^P = \alpha_i^P + \beta_i^P Policy_i + f(quarter_i^P) + \gamma^P \boldsymbol{X}_i^{GP} + \theta^P \boldsymbol{Age}_i^P \\ + \lambda^P Province_i^P \times Birthyear_i^P + \tau_t^P + \nu_i^P \end{aligned} \quad (1)$$

where Y_i^P denotes a parent's outcomes; $Policy_i$ is a dummy variable equal to 1 if the mother's/father's date of birth is from 1980Q1. $f(quarter_i^P)$ is RD polynomials controlling for the distance from the cut-off (1980Q1) in quarters. $\boldsymbol{X}_i^{GP}$ contains pre-determined demographic and social characteristics of grandparents, including their age, literacy, employment status and whether either of them is a member of the communist party[4]. $\boldsymbol{Age}_i^P$ is a non-linear control for the parent's age, including *age* and age^2. We also control for the parent's province-birthyear fixed effects. τ_t^P is interview year fixed effects. Standard errors are clustered by parents' year of birth. The causal effect of the policy on the outcomes of parents is β_i^P. However, our main focus is to examine the spill-over effects of the policy on the second generation: children. The regression estimating the intergenerational effects on children's health outcomes takes the form of:

$$\begin{aligned} Y_i^C &= \alpha_i^C + \beta_i^C Policy_i + f(quarter_i^P) + \gamma^C \boldsymbol{X}_i^{GP} + \theta^C \boldsymbol{Age}_i^P + \delta^C \boldsymbol{X}_i^C \\ &+ \lambda^C Province_i^P \times Birthyear_i^P + \zeta^C Province_i^P \times Birthyear_i^C + \tau_t^C + \nu_i^C \end{aligned} \tag{2}$$

where Y_i^C denotes a child's health outcomes. In addition to the pre-determined characteristics of the grandparents $\boldsymbol{X}_i^{GP}$ and the parent's non-linear age control $\boldsymbol{Age}_i^P$, we also control for the characteristics of children $\boldsymbol{X}_i^C$, in particular, age and gender, and province-by-children's birthyear fixed effects.

In our RD design, $f(quarter_i^P)$ are RD polynomials controlling for the distance from the cut-off in quarters. We use a linear RD polynomial in the baseline specifications (Gelman and Imbens, 2019), and higher orders of RD polynomials in our robustness checks. Additionally, in robustness testing, we will also examine specifications that include the interaction between the treatment variable and the running variable $Policy_i \times f(quarter_i^P)$. This interaction term allows for different functions on either side of the cut-off.

4.2 Identifying Assumptions

Continuity assumption. The first assumption to make the RD design valid is the smoothness of the covariates at the cut-off point. We expect that the changes in our potential outcomes are solely due to the treatment initiated at the cut-off point. No other changes or discontinuities occur at the policy cut-off. Since our treatment is that the parents were born after the cut-off policy date, this assumption is only satisfied when all other relevant covariates related to the parent's birth date, in this case, grandparents' characteristics are continuous in 1980Q1.

We conduct a balance check on a list of predetermined characteristics of both maternal and parental grandparents using the main specifications and controlling for the time trend in the grandparents' birth year, as these younger people are naturally more likely to be treated by the policy. These characteristics include birth year, literacy, employment status and whether either of the grandparents is a member of the communist party. Table 3

[4]Party members were urged to "take the lead" in the one-child policy campaign. In September 1980, the Central Committee of the Chinese Communist Party issued an "Open Letter" to all Party and Youth League members, asking them to lead the way in implementing the policy (Committee, 1984; White, 1990).

and Table 4 show evidence that our design satisfies the continuity assumption. Across all specifications, we do not observe discontinuities of grandparents' pre-determined characteristics at the cut-off quarter.

Table 3. Pre-determined characteristics of mothers' parents

	Dependent variable is:						
	Maternal grandfather			Maternal grandmother			Either
	(1) Age	(2) Literacy	(3) Unemployed	(4) Age	(5) Literacy	(6) Unemployed	(7) Communist
Policy	-0.006	-0.086	0.082	-0.007	0.019	-0.019	-0.014
	(0.028)	(0.062)	(0.043)	(0.018)	(0.027)	(0.088)	(0.016)
Mean	59.233	0.822	0.141	61.094	0.622	0.256	0.154
Observations	3,285	3,206	3,285	3,259	3,240	3,259	3,239
R^2	0.998	0.200	0.165	0.999	0.227	0.133	0.173

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The policy cut-off is the first quarter of 1980. Regressions include mothers born within 11 quarters around the policy cut-off. Columns (1) (2) (3) show the characteristics of mothers' fathers, and columns (4) (5) (6) show the characteristics of mothers' mothers.

Table 4. Pre-determined characteristics of fathers' parents

	Dependent variable is:						
	Paternal grandfather			Paternal grandmother			Either
	(1) Age	(2) Literacy	(3) Unemployed	(4) Age	(5) Literacy	(6) Unemployed	(7) Communist
Policy	0.022	-0.018	0.001	0.041	0.058	-0.032	0.015
	(0.130)	(0.054)	(0.026)	(0.109)	(0.044)	(0.049)	(0.036)
Mean	58.556	0.838	0.141	61.244	0.581	0.221	0.154
Observations	2,259	2,217	2,259	2,235	2,207	2,235	2,228
R^2	0.965	0.306	0.196	0.990	0.300	0.181	0.206

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the fathers' birth years. The policy cut-off is the first quarter of 1980. Regressions include fathers born within 9 quarters around the policy cut-off. Columns (1) (2) (3) show the characteristics of fathers' fathers, and columns (4) (5) (6) show the characteristics of fathers' mothers.

No manipulation. The second assumption ensures that participants are unable to sort themselves on either side of the cut-off point. In our context, parents' birth quarters must not be manipulated around the policy date. Even though some families were pre-aware of the policy in early 1979, because grandmothers need a gestation period of ten months, there is minimal to no opportunity for manipulation at the cut-off quarter (1980Q1).

5 Results

5.1 Baseline Results

We show our baseline results using mothers' birth information in Table 5. Sickness outcomes, which have the most observations, are available for all children in the CFPS data because they include responses from both adult proxies and the children themselves. In contrast, the other three outcomes require self-reporting or children's presence at the interview, resulting in smaller sample sizes.

The children of mothers born after the policy came into effect are less likely to be sick, rate their health status better, and show better overall health at the time of the interview. Specifically, the policy is associated with a 1.8 percentage point (pp) decrease in the likelihood of being sick in the last month, an increase of 8.2pp of children rating themselves as healthy, which represents around a 20% increase compared to the mean, and a 1.8pp increase in the probability that the interviewer rating the child as being in good health. Moreover, these children are also reported to have better mental health, being 8.2pp less likely to be distressed. The results suggest significant improvements in the physical and mental health of children born to mothers after the policy came into effect. In Figure 2, we also present a visual representation of the policy's impact on various child health outcomes using maternal birth information[5]. These figures complement the regression results and show a clear improvement in children's health and mental well-being associated with the policy-taking effects.

Table 5. Children's results using mothers' birth information

	(1) Sick	(2) Self-rated health	(3) Observed health	(4) Distress
Policy	-0.018**	0.082***	0.018***	-0.082**
	(0.005)	(0.018)	(0.004)	(0.024)
Mean	0.263	0.379	0.980	0.123
Observations	3,056	1,176	1,564	1,175
R^2	0.103	0.122	0.069	0.254

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The policy cut-off is 1980Q1. Regressions include children of mothers born within 11 quarters around the policy cut-off. Only those aged 10 and over responded to the self-rated health question, and self-rated health is equal to 1 if children rate themselves as healthy. The scale for interviewer-observed child health ranges from 1 (worst) to 7 (best). The observed health variable equals 1 if the rating is greater than or equal to 4. The interviewers only assessed the health of those children who were present at the interview. Distress in column (4) equals to 1 if Kessler Psychological Distress Scale (K6) is larger than or equal to 5 or CES-D8 is larger than or equal to 7 or CES-D20 is larger than or equal to 16.

[5]See Figure D.1 for quadratic polynomial regressions.

Table 6. Children's results using fathers' birth information

	(1) Sick	(2) Self-rated health	(3) Observed health	(4) Distress
Policy	-0.072**	-0.084	0.046**	-0.042
	(0.025)	(0.045)	(0.014)	(0.036)
Mean	0.283	0.384	0.978	0.102
Observations	2,100	666	957	666
R^2	0.100	0.136	0.102	0.257

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the fathers' birth years. The policy cut-off is 1980Q1. Regressions include children of fathers born within 9 quarters around the policy cut-off.

In Table 6, we show the results for children using fathers' birth information. The results are consistent with those based on mothers' birth information but are less statistically efficient. Children of fathers born after the policy cut-off are less likely to be sick (by 7.2pp) and have better overall observed health (by 4.6pp). They also tend to be less distressed, although this result is not statistically significant. Given the RDD estimation strategy, the results are local around the policy cutoff for the specific selected sample mentioned in Subsection 3.2. Unlike the findings using birth information from mothers, there is no observed improvement in the children's self-rated health. Due to a much smaller sample size and narrower bandwidth of the fathers' data[6], these results tend to be less statistically powerful and should be interpreted with caution.

Our analysis will later on focus mothers' data for several reasons. First, datasets using father information have smaller sample sizes and narrower bandwidth, which reduces our study's statistical power. Second, research indicates that urban Han daughters have particularly benefited from the demographic patterns created by the one-child policy. In particular, many studies show that daughters in urban Han families received more resources and opportunities, leading to improved outcomes in education and overall well-being (Fong, 2002; Huang et al., 2016a, 2021; Veeck et al., 2003; Zhang, 2019). Therefore, we will primarily use the mothers' dataset for our main identification strategy in the following analysis.

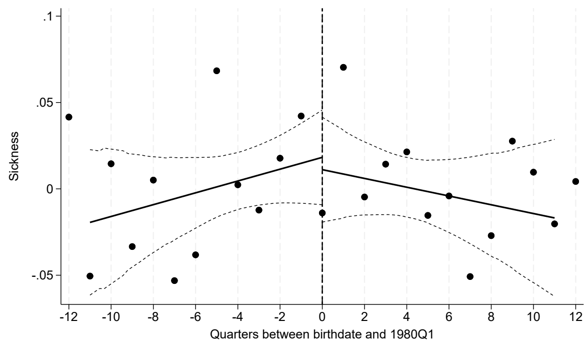
5.2 Robustness Checks

Bandwidth sensitivity. We show bandwidth sensitivity in Appendix Figure B.1. Each sub-graph reports coefficient estimates and 90% confidence intervals of our main results

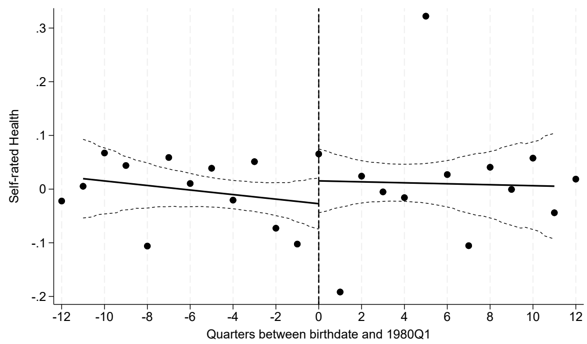
[6]The sex-specific marital status distribution of CFPS is characterized by a higher proportion of unmarried males. The average of the optimal bandwidths for fathers' data is 9 quarters, compared to 11 quarters for mothers' data.

Figure 2. RD plots for all outcomes

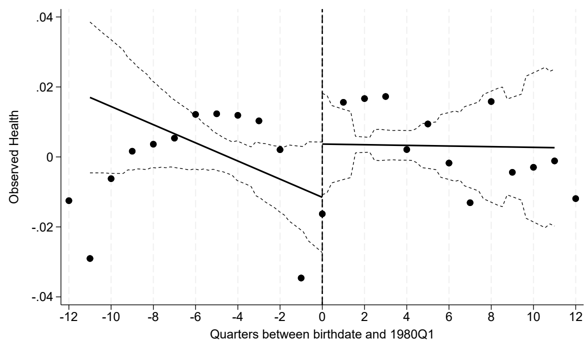
(a) Sick



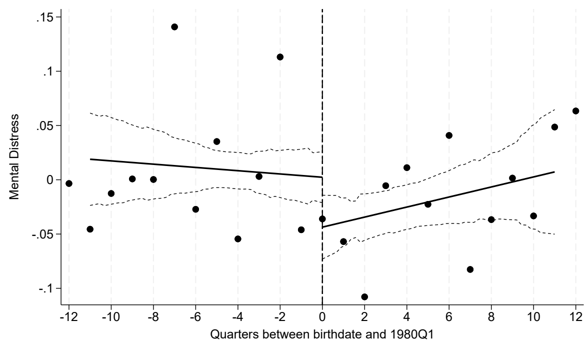
(b) Self-rated health



(c) Observed health



(d) Distress



Notes: The points depict binned residuals from a main regression of the outcome variable on a linear polynomial in birth quarter, along with other control variables. Solid lines display local linear regressions, separately estimated on each side of the cut-off, with dashed lines indicating 90% confidence intervals. Figure D.1 in Appendix D displays RD plots for quadratic polynomial regressions.

for bandwidths ranging from 5 to 15 quarters. Our baseline results show robustness to changes in bandwidth.

For the sickness outcome, the policy effect remains negative and stabilizes at a small, significant level as the bandwidth increases from 10 quarters onwards. For self-rated health status, the point estimates are quite stable and positive around 0.8pp and are statistically significant when the bandwidth is greater than 8 quarters. The estimates for observed health show slight fluctuations but generally remain positive and stable. Lastly, for children's mental health, the policy effects become more pronounced with larger bandwidths, showing a consistent reduction in distress. This may be due to increased statistical power with a larger number of observations and the smaller bandwidth chosen for the main analysis, as shown in Table 2.

Choice of polynomial orders. In our main analysis, we use a linear polynomial of our running variable – mothers' birth quarters, which is the most common choice in RD designs. We show the sensitivity of our results to RD polynomials up to the fourth order in Appendix Figure B.2. The results suggest that our findings are not sensitive to the choice of polynomial order, reinforcing the robustness of our main results. The only exception is the sickness outcome in Appendix Figure B.2a, where policy effects disappear when we move to a third order or fourth order polynomial, although the point estimates still suggest a slightly negative effect. While higher order polynomials may obscure the significance of the sickness results, the general trend remains consistent with our main analysis.

Different specifications. We test the robustness of our main results using various specifications, as shown in Appendix Table B.1. These include models where the treatment is interacted with the running variable and with the quadratic of the running variable, changing the triangular kernel weight to no weights or panel weights provided by the data, and conducting a donut exercise that excludes observations near the cut-off (within 1 quarter from the cut-off). In general, the results are consistent with our main findings although we do lose efficiency in some results.

We first include the linear interaction of the treatment with the running variable in column (1) to allow for differences in slopes on either side of the cut-off when making extrapolation. We see some negative effects on the interaction term for the sickness outcome and the results for other outcomes are consistent with the main. We further add the interaction with the quadratic term of the running variable in column (2) to account for non-linearities in the treatment effects. The point estimates generally get larger when we do so and maintain the consistency of the results.

In columns (3) and (4), we remove the kernel weights and replace them with the panel weights to see if our results are sensitive to the weighting schemes. Here, we lose the significance on observed health status, but the point estimates suggest the same direction. Other outcomes remain stable. In column (5), we conduct a donut exercise by removing all observations close to the policy cut-off – 1 quarter, and keeping the rest of the sample to fit our main specification. The consistent results suggest that the policy effects are not driven only by observations close to the cut-off. In general, we argue that these results support the reliability of our main results across different model specifications.

Placebo tests. Our identification strategy assumes that there is a discontinuity in the treatment at the policy implementation date. One way to test against failures of this assumption is with placebo tests of different policy cut-offs. In Appendix Figure B.3, we reproduce our main specification results with alternative policy cut-offs up to 4 quarters prior and post the policy cut-off employed – 1980 Q1. We find no significant effects from the policy at the placebo cut-offs for sickness and self-rated health. However, for observed health, the effects become stronger as we move the policy cut-off to later quarters. This trend is not concerning because it suggests that the effects seen from 1980 Q1 onwards are driven more strongly by children born to mothers who were born later in the year, as suggested in Figure 2c. For children's mental distress, we observe a significant effect at the 1979 Q4 placebo cut-off. While this effect is noteworthy, it does not persist across other outcomes, suggesting it may be due to random variation rather than a systematic issue with our identification strategy.

5.3 Heterogeneous Effects

Next, we examine the heterogeneous effects of the policy by children's gender and age groups.

Table 7 presents the effects of the policy passed onto boys and girls separately. We run our main specification on subsamples of boys and girls separately. Boys born to a mother born after the policy have significantly higher self-rated health and lower distress, while the effects on sickness and observed health are not significant. For girls, there is a significant improvement in observed health but no significant effects on others. The effects on self-rated health and distress for girls are modest and less pronounced. This difference may be due to the tendency of adolescent girls to rate themselves more conservatively in self-assessments, whether in self-rated health or distress levels (Boerma et al., 2016; Van Droogenbroeck et al., 2018). Research has shown that girls often report lower self-rated health and higher levels of psychological distress than boys, which may explain the observed discrepancies between self-rated and observed health outcomes (Breidablik et al., 2009; Jerdén et al., 2011).

Table 7. Heterogeneous effects by gender

	Sick		Self-rated health		Observed health		Distress	
Gender groups	Boy (1)	Girl (2)	Boy (3)	Girl (4)	Boy (5)	Girl (6)	Boy (7)	Girl (8)
Policy	-0.036	-0.002	0.178**	0.039	0.000	0.031**	-0.108**	-0.032
	(0.034)	(0.031)	(0.057)	(0.073)	(0.010)	(0.011)	(0.039)	(0.044)
Mean	0.26	0.25	0.46	0.43	0.98	0.98	0.14	0.12
Observations	1,601	1,455	599	577	811	753	598	577
R^2	0.127	0.108	0.184	0.250	0.163	0.116	0.303	0.300

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The policy cut-off is 1980Q1. Regressions include children of mothers born within 11 quarters around the policy cut-off.

Table 8 presents the impact of the policy on mothers passed on to children across different age groups (0-12 and 13-15). We group children under 12 as those who are in primary school or less than primary school age, and those aged 13-15 as those in middle school according to the Chinese education system. For the younger age group (0-12), there is a significant improvement in observed health and a reduction in distress for those later in primary school (10-12). For children aged 13-15, we see that they are less likely to be sick and perceive themselves to be in better health. These two health outcomes complement each other, indicating a generally better health status for both age groups.

Table 8. Heterogeneous effects by age groups

	Sick		Self-rated health		Observed health		Distress	
Age groups	≤ 12	> 12	≤ 12	> 12	≤ 12	> 12	≤ 12	> 12
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Policy	-0.009	-0.086*	0.051	0.132***	0.036***	-0.022	-0.131**	-0.025
	(0.013)	(0.035)	(0.048)	(0.028)	(0.006)	(0.033)	(0.040)	(0.014)
Mean	0.29	0.16	0.48	0.41	0.98	0.98	0.13	0.14
Observations	2,565	491	643	533	1,198	366	642	533
R^2	0.102	0.204	0.170	0.167	0.125	0.132	0.250	0.374

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The policy cut-off is 1980Q1. Regressions include children of mothers born within 11 quarters around the policy cut-off.

6 Mechanisms

In this section, we discuss potential mechanisms that we can investigate empirically from the CFPS data. First, better health in children can be explained by higher investment in child health when family sizes become smaller, which is consistent with the quantity-quality trade-off. Second, we provide evidence that parents affected by the policy show improvements in their health status, which can subsequently benefit their offspring. Finally, parenting practices and parent-child interactions reveal that these parents are highly responsive and often put less pressure on their children, explaining the lower levels of psychological distress observed among their children.

Higher investment in child health. The first mechanism is increased investment in children's health, as suggested by the quantity-quality trade-off. We expect children born to parents affected by the policy will receive higher investments from their parents, given the fact that their family size is reduced. Table 9 presents the estimated effects of the policy on family income, overall expenditure, and expenditure specifically allocated towards children's health. We observe a statistically significant increase in total family income, although the magnitude is negligible compared to the mean (column 1)[7]. Higher family income can be attributed to either greater inherited wealth or higher personal

[7]Total family income comprises five components: wage income, total/net business income, property income, transfer income, and other income.

income thanks to better education. In the next analysis, we present evidence showing that policy-affected mothers are more likely to have no siblings, potentially receiving more wealth from their parents, and they also tend to be more educated than those not affected by the policy.

In terms of expenditure, there is also no significant difference in both total expenditure (column 2) and direct medical expenses for children (column 3). Similarly, the likelihood of children having public insurance is the same for those born to policy-affected parents and those not (column 4). However, children born to mothers affected by the policy are 28.2% more likely to have commercial or private health insurance (column 5). Additionally, column 6 shows a considerable increase in parents' spending on children's commercial insurance (0.467, a 33% increase compared to the mean). This empirical evidence suggests that policy-affected mothers are more concerned about their children's health. They allocate more resources towards child healthcare and invest in preventive measures such as health insurance. However, our analysis of fathers' data reveals no increase in investment in children's health (Table D.1), which emphasizes that the improvements in child health primarily come from their mothers' concerns and investments.

Table 9. Family income and expenditure

	(1) Total Income	(2) Total Exp.	(3) Med. Exp.	(4) Public Ins.	(5) Commercial Ins.	(6) Commercial Ins. Spending
Policy	0.221**	0.084*	-0.215	0.019	0.062***	0.467***
	(0.070)	(0.036)	(0.143)	(0.029)	(0.009)	(0.094)
Mean	10.702	10.939	5.370	0.714	0.220	1.412
Observations	3,006	3,004	1,463	3,048	3,035	3,031
R^2	0.277	0.320	0.234	0.164	0.079	0.088

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The first two are taken from family level expenditure and the rest are directly on children. Total family income comprises five components: wage income, total/net business income, property income, transfer income, and other income. We take natural logs of total income and expenditure (columns 1 and 2), medical expenditure (column 5), and commercial insurance spending (column 6). Public and commercial insurance in columns (3) and (4) are binary variables.

Intergenerational transmission effects. The second possible mechanism for improved child health is the intergenerational transmission of health and household characteristics across generations. Existing literature has shown fertility patterns can be transmitted from parents to their offspring (Kolk, 2014). Reduced fertility would lead to increased human capital investment per child, which aligns with our narrative on investing in children's health. Moreover, we anticipate that the policy will have positive causal effects on mothers' health, which can, in turn, be passed on to their children (Emanuel et al., 1992; Eriksson et al., 2005).

Table 10. Policy effects on mothers' demographic characteristics

	(1) No siblings	(2) Number of children	(3) College+
Policy	0.104**	-0.168***	0.137***
	(0.029)	(0.041)	(0.019)
Mean	0.103	1.661	0.228
Observations	3,135	3,146	3,146
R^2	0.292	0.246	0.233

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. No siblings is a dummy variable and College+ is column (3) takes 1 if a mother has a college degree or higher.

Table 10 displays the directly intended effects of the policy. Those mothers affected by the policy are more likely to have no siblings (column 1) and tend to have fewer children (column 2). Their fertility rate decreases by more than 10% compared to the average[8]. Furthermore, they receive better education, with their likelihood of attending college increasing by 13.7 percentage points, which is over 60% above the mean (column 3). This result is consistent with other studies examining the impacts of the one-child policy on women's education (Huang et al., 2016a; Qin et al., 2017). This positive effect on education can also explain higher family incomes we observe above.

Table D.2 presents the policy effects on fathers' demographic characteristics. We observe contrary effects on fathers: their sibling sizes do not change and they tend to have more children. Additionally, we do not see any difference in the rate of attending college, suggesting that the policy does not improve fathers' education achievement.

In terms of their health status, generally, we can see mothers who were born after the policy date have better health outcomes (Table 11). Particularly, for mothers born after the policy date, the likelihood of experiencing physical discomfort in the last two weeks decreases by 8.2 percentage points (column 1) and the probability of being diagnosed with a chronic disease in the past six months also decreases by 6.2 percentage points (column 2). These effects are not only statistically significant but also substantial in magnitude. Columns 3 and 4 present the policy effects on their self-rated health. Although they are more likely to rate their health as inferior (column 3), only 22.96% of the sample think they have good health. Therefore, we further constructed a variable called "unhealthy", which was coded 1 if they perceived their body as very unhealthy. A negative and significant estimate demonstrates that the likelihood of mothers being very

[8]We believe the policy does not influence their decision on the number of children they have. Mothers within our selected bandwidths were born between 1977 and 1982, making them 34 to 39 years old when the policy was eliminated in 2016, which means they could still have another child. Additionally, when the policy was still effective, richer couples in urban areas were willing to pay fines to have another child (Burgess and Zhuang, 2002; Li et al., 2008).

Table 11. Policy effects on mothers' health status

	(1) Discomfort	(2) Chronic Disease	(3) Self-rated health	(4) Unhealthy	(5) Observed health	(6) Distress
Policy	-0.082**	-0.062**	-0.112**	-0.066**	0.004	-0.009
	(0.030)	(0.016)	(0.037)	(0.019)	(0.004)	(0.020)
Mean	0.245	0.074	0.225	0.186	0.981	0.175
Observations	2,477	2,477	3,141	3,141	2,632	2,476
R^2	0.052	0.046	0.219	0.129	0.040	0.150

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. Discomfort takes 1 if a mother reported physical discomfort in the last two weeks. Chronic disease is a dummy variable indicating whether a mother was diagnosed with a chronic disease in the past six months. Self-rated health is a binary variable, with 1 indicating good health, while Unhealthy takes 1 if they rated themselves as very unhealthy. Interviewer-observed health is a binary variable, with 1 indicating good health. Distress is a binary variable where 1 indicates psychological distress.

unhealthy is lower for those affected by the policy (column 4). With regard to interviewer-observed health and mental distress, however, we do not see any significant effects of the policy (column 5). We found similar results when examining the policy effects on fathers' health, with substantial improvements in physical health but insignificant improvements in mental health (Table D.3). These results support our main analysis that children born to policy-affected fathers have better physical health but no improvement in mental health. However, we still interpret these results with caution due to smaller sample sizes and a narrower bandwidth associated with fathers' dataset.

Parenting and Family interactions. Finally, we examine the parenting practices and interactions between parents and children to explain the lower level of distress among children whose parents are affected by the policy. Previous research has extensively examined the relationship between child-rearing practice and children's anxiety. Children may lose their chances to advocate for their interests under parental psychological control, which triggers higher levels of mental distress (Chyung et al., 2022; Luebbe et al., 2014; Rapee, 1997). McCoby (1983), building on the work of Baumrind (1971), identifies four parenting styles characterised by levels of demandingness and responsiveness. Parental demandingness or control significantly influences children's anxiety levels (Pinquart, 2017). High demands from parents cause worry and anxiety, especially for those with executive functioning deficits to manage these concerns. Conversely, low parental demands reduce anxiety among children because they may not be worried about meeting parents' expectations (Lo et al., 2020; Soysa and Weiss, 2014; Wong et al., 2019). Meanwhile, parental responsiveness is another important element that decides the level of anxiety among children. High responsiveness from parents strengthens the family bond and fosters children's social and emotional development, whereas children with less responsive parents are more prone to mental disorders and struggle with social functioning (Davidov and Grusec, 2006; McCoby, 1983; Miller-Slough et al., 2018).

We constructed several variables from the CFPS surveys to explore this mechanism. Table 12 reports the policy effects on the interactions between parents and children. We can see positive and significant impacts on overall home environment, as parents are more actively involved in communicating with their children (column 1). However, there is no significant difference in their concern for children's education (column 2). From children's responses[9], we see children more frequently communicate with their parents, either talking or arguing (columns 3-4). The children born to policy affected mothers may be partly worse off on some socio-emotional dimensions or behaviors within the household suggested by column (3). For those early teenagers who are beginning to seek independence, the increase in arguments may be a natural part of enhanced communication and may not necessarily indicate a worse environment[10]. Though the coefficient is positive, we see null effects of children observing parents quarrelling with each other in column (5). Overall, our results suggest that the conversations within the household increase, and as a result, children also have more opportunities to speak up for themselves.

Table 12. Interaction between parents and children

	Interviewers' observations		Children's Responses		
	(1) Active Communication	(2) Care about Education	(3) Quarrel	(4) Heart-to-heart Talk	(5) Parents Quarrel
Policy	0.020**	0.010	0.788***	0.422	0.349
	(0.005)	(0.009)	(0.187)	(0.450)	(0.266)
Mean	0.869	0.855	1.335	2.527	0.813
Observations	2,530	2,573	1,090	1,022	1,063
R^2	0.572	0.396	0.127	0.127	0.127

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The first two variables are dummy variables, showing interviewers' observations on whether parents communicate with their child actively and on whether home environment indicates parents care about their child's education. The next three variables are only reported by children aged 9-15. Quarrel refers to the number of times children quarrelled with their parents last month (column 3). Heart-to-heart talk refers to the number of times children had a heart-to-heart talk with parents last month (column 4). The last variable (column 5) refers to the number of times the parents quarrelling with each other in the last month.

Meanwhile, the responses of parents in Table 13 suggest that parents born after the policy cut-off tend to be more responsive, but at the same time do not put too much educational pressure on their children. They are 10.5 pp more likely to forgo watching TV to avoid disturbing their children (column 1), and they are 2.7 pp more likely to discuss

[9]Only children aged 9-15 provide answers to these questions.

[10]Table D.6 displays the results when we restrict our sample to those aged 9 or above. The results stay consistent, suggesting more communication between parents and children.

school activities with their children (column 2), which is consistent with the better within household communication results above. They tend to exert not too much pressure on their children, as evidenced by positive but insignificant results of likelihood of checking children's homework or requiring their children to complete homework (column 3). We also see null effects of these parents imposing restrictions on their children's TV watching (column 4). In addition, these parents or households are more likely to have saved for their children's education from an early age, with the effect of increasing the probability by 4.8 pp, 18% compared to the mean (column 5)[11].

Table 13. Parental care

	(1) Give up watching TV	(2) Discuss School	(3) Homework Check	(4) TV Restriction	(5) Save for Education
Policy	0.105***	0.027**	0.032	0.036	0.048***
	(0.024)	(0.010)	(0.019)	(0.023)	(0.010)
Mean	0.257	0.105	0.334	0.194	0.264
Observations	1,942	2,017	1,978	2,000	3,048
R^2	0.099	0.086	0.143	0.088	0.101

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. Outcomes in columns (1) to (4) are dummy variables constructed based on parents' responses: whether parents very often (6-7 times a week) give up watching TV to avoid disturbing their child (column 1), whether parents very often discuss happenings at school with their child this semester (column 2), whether parents very often ask their child to finish homework or check their child's homework (column 3), and whether parents restrict their child from watching TV or restrict the type of TV programs their child could watch (column 4). Column (5) is a dummy variable equal to 1 if the parent answered "yes" to the question of having started saving money for the child's education.

Literature on parenting styles suggests that the ideal parenting style is "authoritative", associated with high responsiveness and an appropriate level of parental control that can promote child autonomy (Baumrind et al., 2010; Doepke and Zilibotti, 2017). We lack sufficient evidence to determine the exact level of control these parents exert on their children; however, our results indicate that policy-affected parents exhibit high responsiveness and put not too much pressure on their children. Children benefit from increased parents' warmth and support, enjoy a more relaxed environment, and have greater freedom in their actions. Our findings align with existing research on Chinese parents, which indicates that increased parental responsiveness towards their children results in lower psychological distress, especially among single kids (Liu et al., 2010; Lu and Chang, 2013).

[11]We see a general null effects on parental practices and parent-child interaction from the fathers' side in Table D.4 and Table D.5. In general, these results support our main finding using fathers' dataset that there is no discontinuity in children's level of distress at the policy cut-off.

7 Conclusion

China's one-child policy, although formally abolished and relaxed in 2016, has had long-lasting and profound effects on the entire population. For 35 years, the policy restricted most Chinese families to one child, directly affecting at least two generations and leading to many unexpected consequences for family structure. In this paper, we empirically examine the intergenerational effects of the one-child policy on the health outcomes of children whose parents were directly affected by it. Using data from the China Family Panel Studies (CFPS), we provide causal evidence on how the policy, which significantly altered family size and dynamics, affected the physical and mental health of the next generation.

Our results show that children of mothers born after the policy was implemented show significant improvements in both physical and mental health. These children are less likely to be ill, rate their health more positively, and are observed by interviewers to be healthier. They also show lower levels of psychological distress, suggesting a positive impact of the policy on their mental well-being.

We emphasize the importance of focusing on mothers' side because of the unique benefits observed for urban Han daughters under the one-child policy. These findings are consistent with the theoretical framework of the quantity-quality trade-off, where reduced family size leads to increased investment in the health and education of each child. Furthermore, our results suggest that the intergenerational transmission of health and improved parental health are key mechanisms driving these results. Parents who benefit from the policy tend to be more responsive and less demanding, contributing to the lower levels of distress observed in their children.

Our study contributes to the literature on quantity-quality trade-offs and intergenerational effects by providing comprehensive evidence from a nationally representative dataset. We highlight the broader impact of family planning policies on child health and well-being, and emphasize the need to consider both direct and spillover effects in policy evaluations.

In conclusion, despite its controversial nature and significant social consequences, the one-child policy in China has led to remarkable improvements in the health of the next generation. These findings provide valuable insights for policymakers and researchers interested in the long-term effects of family planning policies and their role in shaping population health and family dynamics.

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Appendix

A Description of Psychological Scales

CFPS uses different scales of mental health distress in different survey waves. One such indicator is the Kessler Psychological Distress Scale (K6), developed by Kessler et al. (2002), which was asked in the 2010 and 2014 surveys. Respondents reported their experiences in the past month on items in Table A.1. We reverse code each item to score as 0 (never), 1 (once a month), 2 (2-3 times a month), 3 (2-3 times a week) and 4 (Almost every day) and aggregate them to a final score ranging from 0 to 24, with higher scores indicting greater depressive symptoms. While a score of 13 usually defines serious mental illness (Kessler et al., 2003), we use a lower threshold of K6 ≥ 5 to indicate moderate mental distress (Prochaska et al., 2012).

Another mental health indicator used in CFPS is the Center for Epidemiologic Studies Depression Scale (CES-D) (Radloff, 1977). The full 20-item CES-D was included in the 2012 and 2016 surveys, while an 8-item version was asked in the 2018 and 2020 surveys. Respondents rated their past-week status on items in Table A.1. Each item was scored as 1 (never(less than one day)), 2 (sometimes (1-2 days)), 3 (often(3-4 days)), and 4 (most of the time (5-7 days)). We reverse code items 4, 8, 12, 16 and aggregate those items, with the 20-item version ranging from 0 to 60 and the 8-item version from 0 to 24. Higher scores indicate more sever depression. The CES-D20 categorizes scores as follows: ≤ 16 indicates no to mild depression, 17-23 indicates moderate depression, and ≥ 24 indicates severe depression (Bi et al., 2023). In this paper, we use a used CES-D20 cut-off of 16, corresponding to an CES-D8 cut-off score of 7, as these scores effectively identify individuals at risk of clinical depression in the Chinese context (Bi et al., 2023).

Both K6 and CES-D are frequently used to evaluate psychological distress and serious mental illness (Kessler et al., 2003; Kim et al., 2016; Weissman et al., 1977). Because these scales do not appear in all survey waves, we construct a consistent variable called "distress". A child is coded as 1 (distressed) if their K6 score is ≥ 5 (Prochaska et al., 2012), CES-D8 score is ≥ 7, or CES-D20 score is ≥ 16 (Bi et al., 2023).

Table A.1. Items of psychological scales

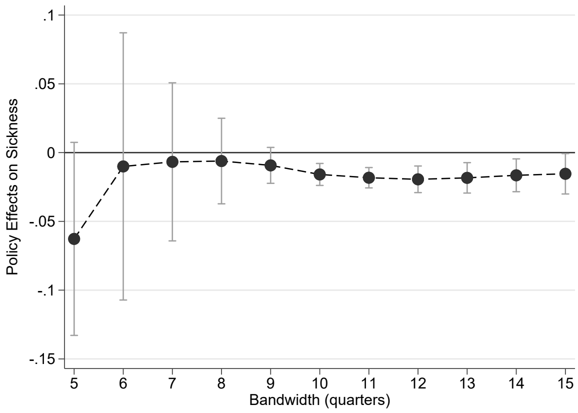
	2010	2012	2014	2016	2018	2020
K6: Please select according to your statuses in the past month.						
(1) Feel depressed and cannot cheer up.	X		X			
(2) Feel nervous.	X		X			
(3) Feel agitated or upset and cannot remain calm.	X		X			
(4) Feel hopeless about the future.	X		X			
(5) Feel that everything is difficult.	X		X			
(6) Think life is meaningless.	X		X			
CES-D: Please select according to your statuses in the past week.						
(1) I am worried about some trivial things.		X		X		
(2) I have a poor appetite and do not want to eat.		X		X		
(3) I feel depressed despite the help from relatives and friends.		X		X		
(4) I find myself not worse than others.		X		X		
(5) I cannot concentrate on things.		X		X		
(6) I am in a low spirit.		X		X	X	X
(7) I find it difficult to do anything.		X		X	X	X
(8) I find the future promising.		X		X		
(9) I feel that I have been a loser for a long time.		X		X		
(10) I feel scared.		X		X		
(11) I cannot sleep well.		X		X	X	X
(12) I feel happy.		X		X	X	X
(13) I talk less than usual.		X		X		
(14) I feel lonely.		X		X	X	X
(15) I find that people are not friendly to me.		X		X		
(16) I have a happy life.		X		X	X	X
(17) I cried or I want to cry.		X		X		
(18) I feel sad.		X		X	X	X
(19) I find that others do not like me.		X		X		
(20) I feel that I cannot continue with my life.		X		X	X	X
Number of items	6	20	6	20	8	8

Notes: This table presents detailed items of the K6 scale and the CES-D scale and in which wave they were elicited.

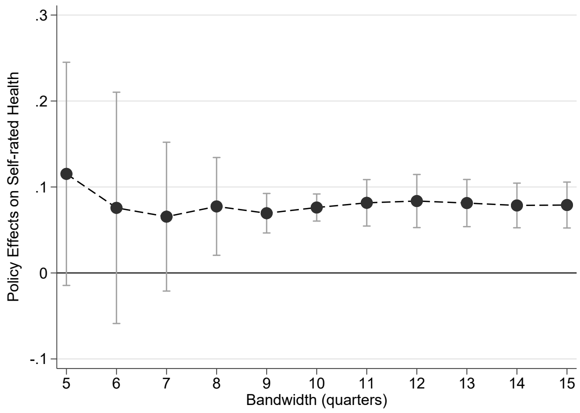
B Robustness Checks

Figure B.1. Sensitivity of results to bandwidth choices

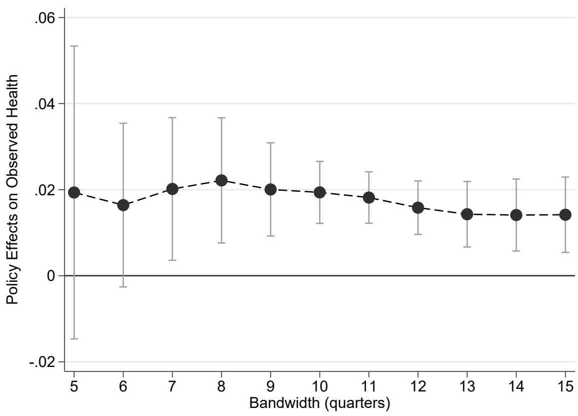
(a) Sick



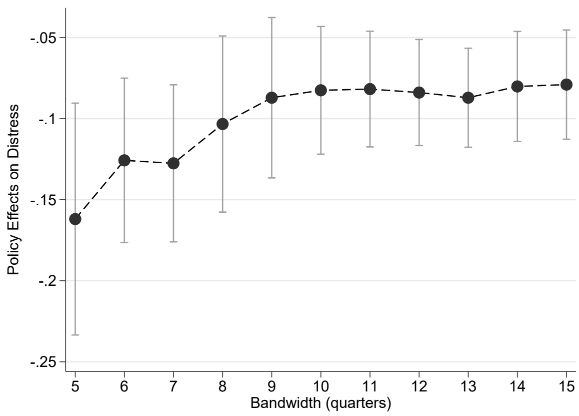
(b) Self-rated health



(c) Observed health

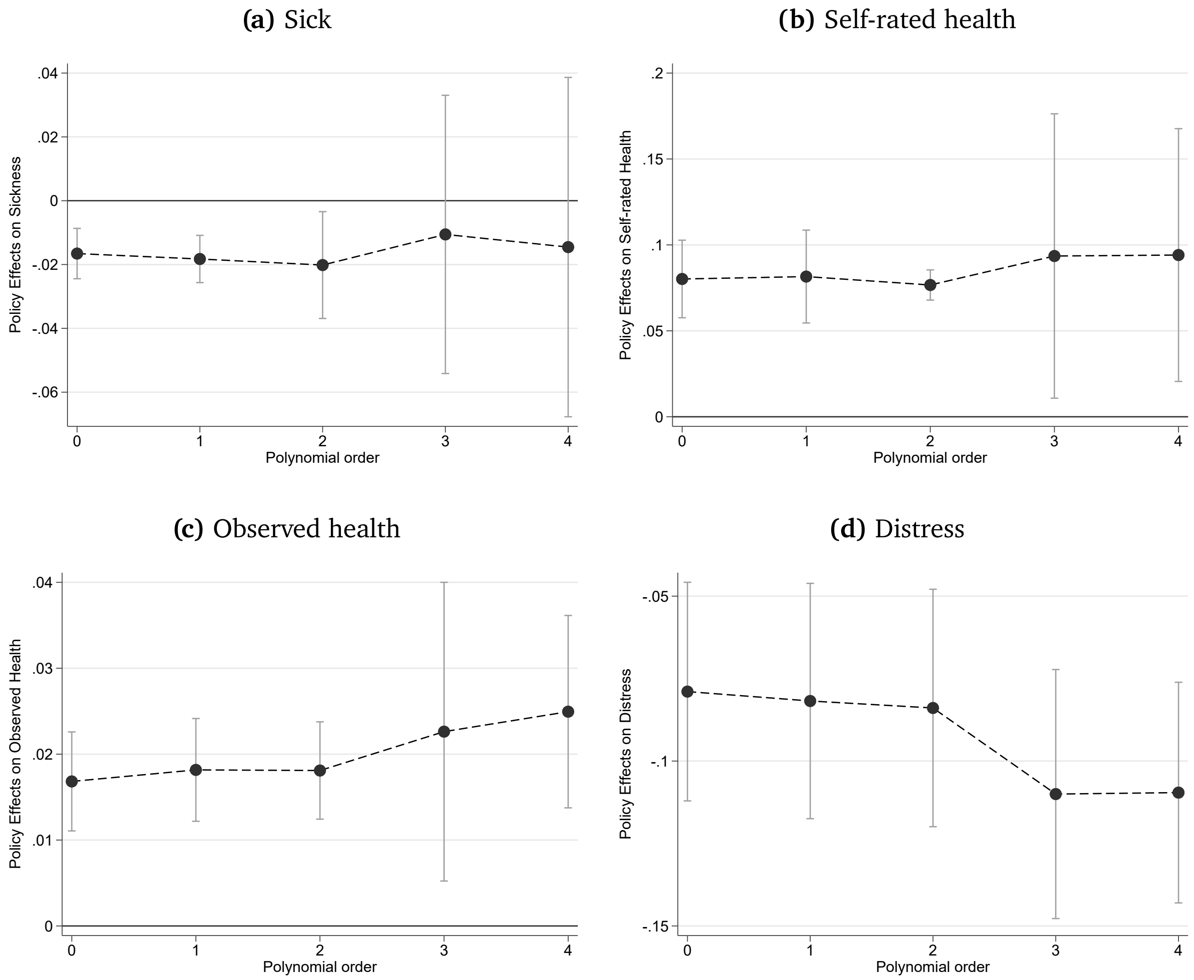


(d) Distress



Notes: Each sub-graph reports coefficient estimates and confidence intervals for different bandwidths from 5 to 15 quarters. Each dot indicates the RD estimate using the specified bandwidth. Capped spikes represent 90% confidence intervals of the estimates.

Figure B.2. Sensitivity of results to different orders of polynomial



Notes: Each dot represents the RD estimate using the specified order of RD polynomial. Capped spikes represent 90% confidence intervals of the estimates.

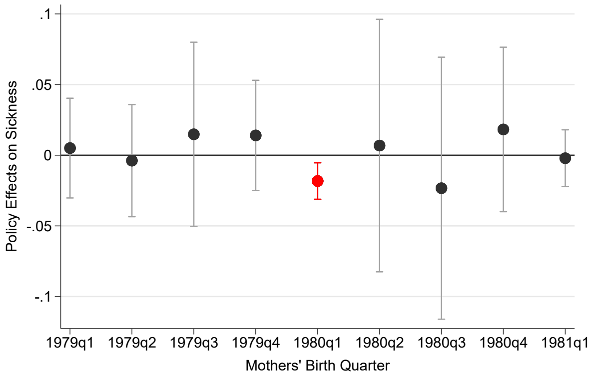
Table B.1. Robustness to different specifications

	Linear Interaction (1)	Quadratic Interaction (2)	No weights (3)	Panel weights (4)	Donut 1Q (5)
Panel A. Dependent variable is: Sick					
Policy	-0.022	-0.018	-0.023	0.014	-0.006
	(0.012)	(0.048)	(0.021)	(0.012)	(0.021)
Policy × Running quarters	-0.006**	-0.005			
	(0.002)	(0.016)			
Policy × Running quarters^2		0.000			
		(0.003)			
Mean	0.263	0.263	0.263	0.253	0.263
Observations	3,056	3,056	3,274	2,907	2,938
R^2	0.104	0.104	0.089	0.106	0.101
Panel B. Dependent variable is: Self-rated health					
Policy	0.079***	0.143*	0.091**	0.193**	0.057
	(0.013)	(0.057)	(0.032)	(0.073)	(0.046)
Policy × Running quarters	0.013	0.059			
	(0.008)	(0.033)			
Policy × Running quarters^2		0.003			
		(0.005)			
Mean	0.379	0.379	0.379	0.378	0.378
Observations	1,176	1,176	1,266	1,251	1,130
R^2	0.123	0.124	0.114	0.211	0.145
Panel B. Dependent variable is: Observed health					
Policy	0.019**	0.043***	0.006	0.009	0.041***
	(0.006)	(0.010)	(0.011)	(0.009)	(0.010)
Policy × Running quarters	0.003	0.031**			
	(0.002)	(0.009)			
Policy × Running quarters^2		0.001			
		(0.001)			
Mean	0.980	0.980	0.980	0.983	0.980
Observations	1,564	1,564	1,666	1,529	1,500
R^2	0.070	0.075	0.068	0.167	0.081
Panel B. Dependent variable is: Distress					
Policy	-0.083**	-0.103***	-0.089**	-0.217**	-0.078*
	(0.023)	(0.020)	(0.023)	(0.067)	(0.033)
Policy × Running quarters	0.005	0.021*			
	(0.005)	(0.010)			
Policy × Running quarters^2		-0.002			
		(0.001)			
Mean	0.123	0.123	0.123	0.125	0.125
Observations	1,175	1,175	1,265	1,250	1,129
R^2	0.254	0.255	0.237	0.251	0.263

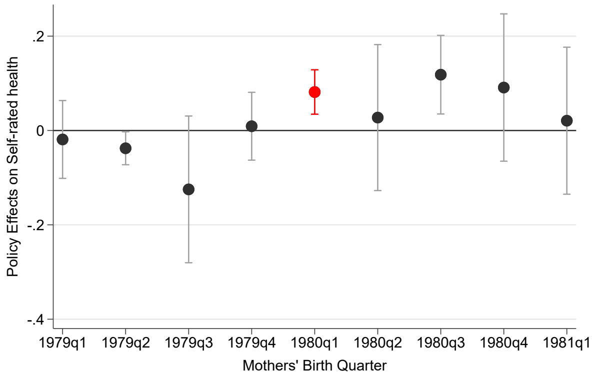
Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The policy cut-off is 1980Q1. Regressions include children of mothers born within 11 quarters around the policy cut-off. CFPS panel weights are used in column (4).

Figure B.3. Placebo 1980 Q1 cut-offs

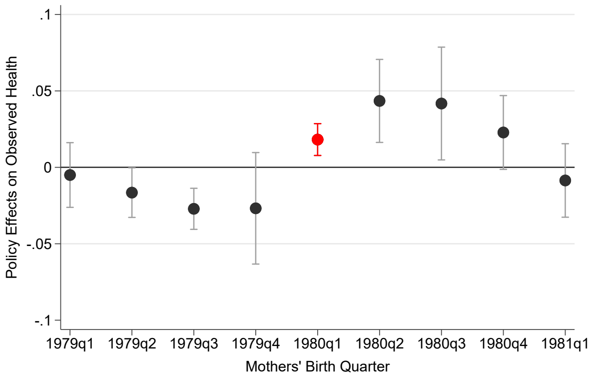
(a) Sick



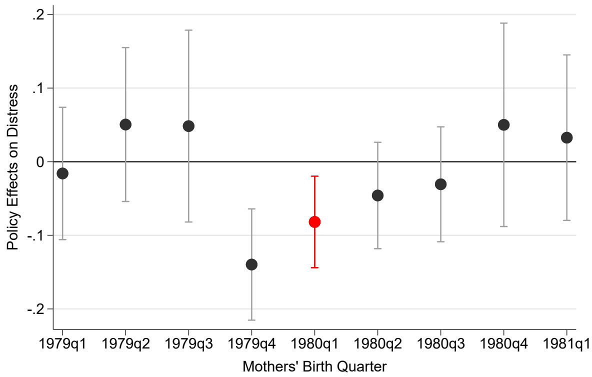
(b) Self-rated health



(c) Observed health



(d) Distress



Notes: This figure tests different policy cut-offs up to 4 quarters prior and post the policy cut-off employed in this paper, at a 1-quarter frequency. The policy cut-off we choose for this paper is 1980Q1, marked in red.

C Results on child growth

Growth indicators. In addition to the main results, we also look at children's growth indicators, which are often used for younger kids to indicate their general health. We use the same specification as our main analysis but exclude controls for the children's gender and age, as the Z-scores already take these factors into account.

Table C.1 shows the effects of the policy on various indicators of child growth. The results generally show null effects on child growth indicators, except for a slightly significant increase in the probability of being overweight, as derived from the BMI Z-score. These additional results suggest that children born to mothers who were born right before or after the policy cut-off have similar growth patterns, except for a marginally higher probability of being overweight. This is generally consistent with some papers that also find null effects of sibling size on children's height and BMI (Zhong, 2014).

Table C.1. Growth indicators for children

	(1) Height-for-age Z-score	(2) Weight-for-age Z-score	(3) Body Mass Index Z-score	(4) Overweight	(5) Obese
Policy	-0.062	0.204	0.175	0.070*	0.004
	(0.039)	(0.111)	(0.157)	(0.034)	(0.015)
Mean	0.138	0.241	0.242	0.094	0.062
Observations	2,801	2,047	2,832	2,832	2,832
R^2	0.149	0.140	0.099	0.051	0.095

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. Z-scores are generated using the WHO Child Growth Standards, accounting for age and gender. Overweight is defined as a BMI-for-age Z-score above 2, while obese is defined as a BMI Z-score above 3.

Weight categories for mothers. Table C.2 shows the effects of the policy on the body mass index (BMI) and different weight categories for mothers. The policy has no significant effect on the BMI of mothers in general. However, we see a small but significant increase in the probability of being in the overweight category for mothers born after the policy cut-off date. This is consistent with the literature looking at the effects of the policy on health in middle age (Islam and Smyth, 2015; Wu and Li, 2012). Meanwhile, we see a reduction in the probability of being obese for these mothers, and null effects on the probability of being in the healthy weight or underweight categories.

These results suggest that the policy had a noticeable effect on the weight distribution of mothers, specifically by increasing the likelihood of being overweight. A possible explanation for the increase in the overweight category but not in obesity could be that the policy led to improved economic conditions and access to food and nutrition, which caused mothers to gain weight and move from a healthy weight to overweight. However, the same improvements in economic conditions and access to health care may have prevented the extreme weight gain that leads to obesity, explaining the decline in obesity rates. This shift in weight categories reflects nuanced changes in maternal health outcomes influenced by policy, highlighting the complex interplay between fertility policy and health behaviors. This effect on mothers' weight outcomes could also be transmitted to their children, as shown in the above Appendix Table C.1.

Table C.2. Weight categories for mothers

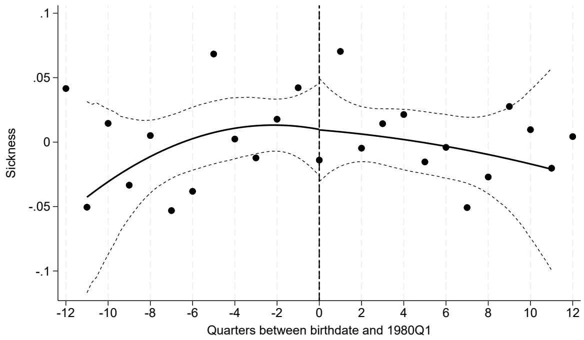
	(1) Body Mass Index	(2) Healthy weight	(3) Overweight	(4) Obese	(5) Underweight
Policy	0.005	-0.056	0.073***	-0.032*	0.015
	(0.119)	(0.030)	(0.016)	(0.014)	(0.023)
Mean	22.529	0.631	0.226	0.065	0.078
Observations	2,612	2,612	2,612	2,612	2,612
R^2	0.197	0.078	0.095	0.088	0.075

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. Healthy weight is defined as a BMI larger than or equal to 18.5 but smaller than 24. Overweight is defined as a BMI above 24, while obese is defined as a BMI above 28.

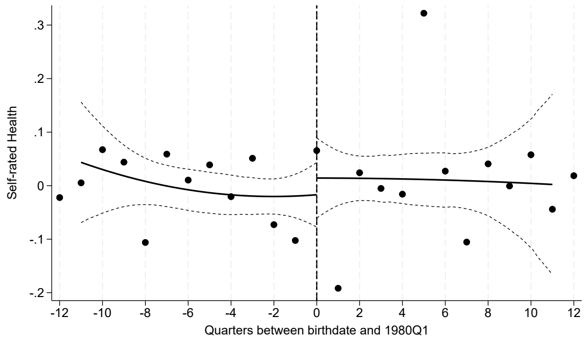
D Additional tables and figures

Figure D.1. Quadratic polynomial: RD plots for all outcomes

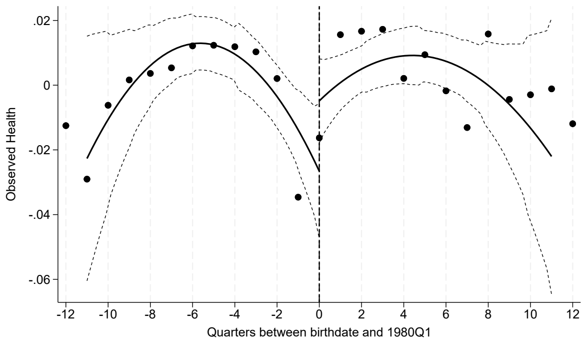
(a) Sick



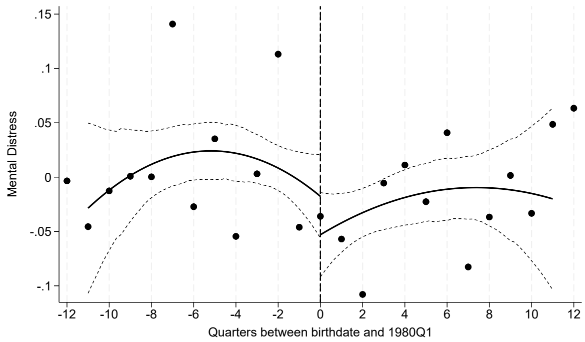
(b) Self-rated health



(c) Observed health



(d) Distress



Notes: The points depict binned residuals from a main regression of the outcome variable on a quadratic polynomial in birth quarter, along with other control variables. Solid lines display quadratic polynomial regression fit, separately estimated on each side of the cut-off, with dashed lines indicating 90% confidence intervals.

Table D.1. Effects from fathers' side: Family income and expenditure

	(1)	(2)	(3)	(4)	(5)	(6)
	Total Income	Total Exp.	Med. Exp.	Public Ins.	Commercial Ins.	Commercial Ins. Spending
Policy	-0.211***	0.017	-0.245	-0.032*	-0.086**	-0.556*
	(0.036)	(0.076)	(0.135)	(0.013)	(0.027)	(0.229)
Mean	10.726	10.943	5.542	0.721	0.206	1.315
Observations	2,542	2,491	1,149	2,533	2,530	2,528
R^2	0.333	0.329	0.222	0.166	0.098	0.106

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the fathers' birth years. The first two are taken from family level expenditure and the rest are directly on children.

Table D.2. Policy effects on fathers' demographic characteristics

	(1) No siblings	(2) Number of children	(3) College+
Policy	-0.048	0.107**	0.037
	(0.032)	(0.033)	(0.089)
Mean	0.215	1.751	0.220
Observations	2,147	2,158	2,158
R^2	0.331	0.327	0.280

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the fathers' birth years. No siblings is a dummy variable and College+ is column (3) takes 1 if a mother has a college degree or higher.

Table D.3. Policy effects on fathers' health status

	(1) Discomfort	(2) Chronic Disease	(3) Self-rated health	(4) Unhealthy	(5) Observed health	(6) Distress
Policy	-0.102***	-0.069	0.124***	0.012	-0.000	-0.021
	(0.017)	(0.047)	(0.022)	(0.025)	(0.008)	(0.036)
Mean	0.188	0.095	0.242	0.163	0.983	0.161
Observations	1,723	1,720	2,153	2,153	1,752	1,723
R^2	0.092	0.077	0.224	0.116	0.056	0.150

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the fathers' birth years. Discomfort takes 1 if a father reported physical discomfort in the last two weeks. Chronic disease is a dummy variable indicating whether a father was diagnosed with a chronic disease in the past six months. Self-rated health is a binary variable, with 1 indicating good health, while Unhealthy takes 1 if they rated themselves as very unhealthy. Interviewer-observed health is a binary variable, with 1 indicating good health. Distress is a binary variable where 1 indicates psychological distress.

Table D.4. Effects from fathers' side: Interaction between parents and children

	Interviewers' observations		Children's Responses		
	(1) Active Communication	(2) Care about Education	(3) Quarrel	(4) Heart-to-heart Talk	(5) Parents Quarrel
Policy	-0.007	0.015	-0.287	-0.432	0.145
	(0.018)	(0.017)	(0.165)	(0.520)	(0.373)
Mean	0.843	0.843	1.140	2.506	0.811
Observations	1,729	1,774	626	609	608
R^2	0.550	0.400	0.168	0.129	0.158

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The first two variables are dummy variables, showing interviewers' observations on whether parents communicate with their child actively and on whether home environment indicates parents care about their child's education. The next three variables are only reported by children aged 9-15. Quarrel refers to the number of times children quarrelled with their parents last month (column 3). Heart-to-heart talk refers to the number of times children had a heart-to-heart talk with parents last month (column 4). The last variable (column 5) refers to the number of times the parents quarrelling with each other in the last month.

Table D.5. Effects from fathers' side: Parental care

	(1) Give up watching TV	(2) Discuss School	(3) Homework Check	(4) TV Restriction	(5) Save for Education
Policy	-0.089	-0.009	-0.037	0.078	-0.097**
	(0.048)	(0.011)	(0.042)	(0.048)	(0.030)
Mean	0.246	0.103	0.304	0.199	0.233
Observations	1,253	1,295	1,259	1,289	2,090
R^2	0.134	0.111	0.175	0.085	0.119

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. Outcomes in columns (1) to (4) are dummy variables constructed based on parents' responses: whether parents very often (6-7 times a week) give up watching TV to avoid disturbing their child (column 1), whether parents very often discuss happenings at school with their child this semester (column 2), whether parents very often ask their child to finish homework or check their child's homework (column 3), and whether parents restrict their child from watching TV or restrict the type of TV programs their child could watch (column 4). Column (5) is a dummy variable equal to 1 if the parent answered "yes" to the question of having started saving money for the child's education.

Table D.6. Effects from mothers' side: Interaction between parents and children, restricted to children above 9 years old

	Interviewers' observations		Children's Responses		
	(1) Active Communication	(2) Care about Education	(3) Quarrel	(4) Heart-to-heart Talk	(5) Parents Quarrel
Policy	0.053***	0.023	0.788***	0.422	0.349
	(0.010)	(0.024)	(0.187)	(0.450)	(0.266)
Mean	0.803	0.801	1.335	2.527	0.813
Observations	1,088	1,114	1,090	1,022	1,063
R^2	0.598	0.425	0.127	0.127	0.127

Notes: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Standard errors are in parentheses and clustered at the mothers' birth years. The first two variables are dummy variables, showing interviewers' observations on whether parents communicate with their child actively and on whether home environment indicates parents care about their child's education. The next three variables are only reported by children aged 9-15. Quarrel refers to the number of times children quarrelled with their parents last month (column 3). Heart-to-heart talk refers to the number of times children had a heart-to-heart talk with parents last month (column 4). The last variable (column 5) refers to the number of times the parents quarrelling with each other in the last month.