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A Bit of Salt, A Trace of Life: Gender Norms and The Impact of a Salt Iodization Program on Human Capital Formation of School Aged Children

Discussion Paper no. 2021-01

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Keywords: Iodine, parental investments, gender attitudes, cognitive skills, non-cognitive skills

JEL Classification: I15, J16, J24, O15

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A Bit of Salt, A Trace of Life:

Gender Norms and The Impact of a Salt Iodization Program on Human Capital Formation of School Aged Children*

Zichen Deng† Maarten Lindeboom‡

Current version: Tuesday 8th September, 2020

Abstract

This paper examines the effects of a massive salt iodization program on human capital formation of school-aged children in China. Exploiting province and time variation, we find a strong positive impact on cognition for girls and no effects for boys. For non-cognitive skills, we find the opposite. We show in a simple model of parental investment that gender preferences can explain our findings. Analyses exploiting within the province, village-level variation in gender attitudes confirm the importance of parental gender preferences. Consequently, large scale programs can have positive (and possibly) unintended effects on gender equality in societies with son preference.

KeyVerbals: Iodine, parental investments, gender attitudes, cognitive skills, non-cognitive skills.

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*The main title is borrowed from Endeavors. Thanks to Douglas Almond, Prashant Bharadwaj, Arnaud Chevalier, Hans Grönqvist, David Johnston, Johannes Kunz, Travis Lybbert, Peter Nilsson, Giacomo Pignataro, Hitoshi Shigeoka and Gerard van den Berg, for helpful conversations and seminar/workshop audiences at RES Symposium, Essen, Toulouse, NEUDC, Kyoto, Milan, Tilburg, Hamburg, Uppsala, and Oslo for useful comments. All errors are our own.

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1 Introduction

Iodine deficiency early in pregnancy can have significant, irreversible effects on the brain development of the fetus (Cao et al., 1994) and can, therefore, have important consequences for human capital formation of children and subsequent socioeconomic outcomes. Indeed, the large medical literature (see Zimmermann, 2011, for a systematic literature review) has found adverse effects of iodine deficiency on later life growth and development, in particular when exposed early in life. Using historical data, recent economics papers such as Feyrer, Politi, and Weil (2017); Adhvaryu et al. (2020) found that cohorts exposed to higher levels of iodine intake have higher labor force participation rates, higher incomes and also higher probabilities of entering top-tier occupations with higher cognitive demands. Although the benefits of micro-nutrient improvement have been documented for decades, still only around 66% of households have access to iodized salt globally.[1] Further, little is known about pathways between early in life exposure and adult productivity and earnings. This impedes our understanding of how developmental trajectories unfold over the life course and who benefits most from large scale interventions.

Our contribution to the extent of literature is to consider and account explicitly for possible interactions with gender preferences in receiving parental investments early in life. These differences are deeply rooted in gender norms, which are important in East and South-East Asia and the Middle East and North Africa. The idea that gender preferences play a role in the western world cannot be excluded either. Biased gender norms may already have an impact on children very early in life and might explain heterogeneous gender effects found in the literature on the long-run effects of early life shocks (Field, Robles, and Torero, 2009; Maccini and Yang, 2009; Adhvaryu, Bednar, Molina, Nguyen, and Nyshadham, 2020). This paper examines the effect of a massive, nationally implemented salt iodization program in China. Evidence of differential program effects by gender would not only provide new interpretation to gender differences in the impact of early in life shocks found in the literature but also shed light on predictions of theoretical models of human capital formation that allow parents to make compensatory and reinforcing investments in different dimensions of human capital. Parental child preferences play an important role in such investments. If compensatory investments of parents are relevant, then large scale interventions may crowd out (or reinforce) private investments and may also affect skills dimensions initially not aimed. Public policy can, therefore, have positive (and possibly) unintended effects on

[1] In 2007, an estimated 31.5% of school-age children (266 million) had insufficient iodine intake. In the general population, this amounted to 2 billion people. See also: WHO, www.who.int/nutrition/publications/micronutrients/FNBvol29N3sep08.pdf, UNICEF, www.who.int/nutrition/topics/idd/en/.

gender equality in societies where gender norms are important. Understanding mechanisms underlying program effects are crucial as any intervention would be blind without knowledge of the mechanisms underlying behavioral responses.

To fight against iodine deficiency-related diseases around the country, the Chinese government implemented a national program of regulating salt to contain iodine in October 1994. At the same time, biennial province-based monitoring was introduced to record the use and iodine content of household salt, along with urinary iodine concentrations among schoolchildren. After the introduction of the program, the urinary iodine concentration reached satisfactory levels from 1995 onward, and the percentage of children who had goiter[2] dropped rapidly. Given the importance of iodine during the gestational period for brain development (Cao et al., 1994; Zimmermann, 2011), we, in the first instance, focus on the potential impact of this policy on cognitive ability and school attainment of children who are affected in utero.

For this, we link iodine deficiency information across locations collected at the start of the intervention to nationally representative rural samples drawn from the *China Family Panel Studies* (CFPS). A unique aspect of the CFPS survey is that math and vocabulary ability from standardized tests were collected besides information on educational attainment and schooling. Similar to Shah and Steinberg (2017), our human capital measurements in the CFPS have the advantage that the same questions were given to each individual in the survey, no matter whether he/she is currently enrolled in school or not. With this data set, our empirical design does not suffer from selection bias caused by censoring individuals who had already left school. To identify the long-term benefits of the salt iodizing policy, we use the national salt iodizing program as a quasi-experiment and exploit geographic variation in goiter prevalence before the intervention. So, we essentially compare improvements in math and vocabulary ability as well as educational attainment and years of schooling of cohorts conceived before and after the salt iodization in areas with varying pre-intervention goiter prevalence. Additionally, we make sure that we are comparing outcome variable trends (by birth cohort) across high and low goiter province of birth in their deviation from each region's average trend. We also include an extensive set of controls, such as pre-treatment province characteristics interacted with cohort dummies to flexibly control for confounding factors at the province level that might deferentially affect cohorts. Our difference-in-differences estimates show that the salt iodization policy has substantial and significant effects on cognition for girls. A one standard deviation (12%) decrease in the

[2] Iodine is an essential component of hormones produced by the thyroid gland. Iodine deficiency can lead to an enlarged thyroid gland located at the base of the neck, which is the most visible consequence of iodine deficiency. Goiter in adulthood does not have severe consequences, while iodine deficiency in utero can lead to impaired neurodevelopment of the child and post-birth cognitive disabilities.

pre-intervention regional goiter rate is associated with an increase of 15% in female math and vocabulary test scores. We also see significant increases in educational attainment and the schooling of women. Yet, for men, we find much smaller and insignificant effects. The improvement in human capital translates to an about 5% increase in income for females.[3]

Gender preferences are important in China. We present a model of human capital accumulation and parental investments as one potential way to rationalize the empirical results described above. In this model, gender preferences play a part, and parental investments interact with different endowments at birth. The model suggests that with preferences on boys, parents tend to compensate more for boys than girls when they are disadvantaged at birth. The model's predictions are in line with our finding of strong and sizable effects for girls and small and insignificant effects for boys. On the other hand, the model suggests that parental investments can be diverted to different dimensions of skills, such as non-cognitive skills, when there is no need to compensate for the cognitive disadvantage at birth. Similarly, with preferences on boys, the diversion will be stronger for boys than for girls. Indeed, we find that the program has positive effects for boys on non-cognitive skill measures. We find no impact on girls in these non-cognitive skill dimensions. Similar to Dahl, Kotsadam, and Rooth (2017); Dhar, Jain, and Jayachandran (2018); Dossi et al. (2019), we proxy gender preferences by gender attitudes, specifically, about the appropriate roles and rights of women and girls. Across all outcomes, we find that the gains in cognition are most significant for girls born in regions with the strongest son preferences. For boys, we see an opposite heterogeneous treatment effects of iodized salt across districts with different gender attitudes. All the evidence above suggests that gender preferences are an indispensable pathway to explaining the gender difference in the program effects.

Our study contributes to at least four strands in the literature. Firstly, we add to the literature on the long-term effects of early-life conditions. Much of early "fetal origins" work (see, among others, Almond, 2006; Van den Berg, Lindeboom, and Portrait, 2006) has focused on demonstrating the impact of extreme, traumatic experiences (disease outbreak, recessions, famines, severe environmental shocks, etc.) in early life. Recent studies (see Niemesh, 2015; Hoynes, Schanzenbach, and Almond, 2016; Feyrer, Politi, and Weil, 2017; Brown, Kowalski, and Lurie, 2018; Adhvaryu et al., 2020) have shifted the focus to estimating gains to exposure to a purposeful large-scale distribution of resources. The nationally implemented intervention in China started just after the launch of the 1993 WHO campaign. The program is, to the best of our knowledge, the largest of its kind. It is also a commonplace, moderate intervention, and has, therefore, relevant external validity—this aids policymakers

[3]To calculate the increase in income, we use the results in Wang (2013), who finds that one year of additional schooling raises income with 15% in China.

in optimizing similar policies in the future. Furthermore, while there exist a large body of studies that look at the impact of in utero exposure to the quantity of food (see Lumey, Stein, and Susser, 2011, for an excellent review of the famine literature), only a few studies (e.g. Field, Robles, and Torero, 2009; Feyrer, Politi, and Weil, 2017; Adhvaryu et al., 2020) have looked at the long-run effects of food *quality* or nutrient intake.

Secondly, and related to the above, we contribute to the discussion on intermediate proxy indicators of long-term outcomes. Adhvaryu et al. (2020); Feyrer, Politi, and Weil (2017) examines the long term effect of a salt iodization program, promoted by a private firm, on lifetime income at later ages. Adhvaryu et al. (2020) find an about 10% income increase for those exposed to the iodine program. We look at the effect of a public program on childhood cognition and education and find substantial effects. Our study thus adds to a full picture of how early-life disadvantage unfolds over the life course. We use measurements of human capital that include standardized numeracy tests for all children, as opposed to most of the previous literature, which only focuses on school enrollment. A few recent studies such as (Figlio et al., 2014; Almond, Mazumder, and van Ewijk, 2015; Bharadwaj, Lundborg, and Rooth, 2017; Shah and Steinberg, 2017) take a similar approach as we do by examining the effects of events in early childhood on cognitive test outcomes during the school years. Most studies use administrative data of developed countries, where standardized tests cover most of the school-going children at a certain age. In developing countries, however, a substantial share of the children is already out of school at young ages, and therefore, a similar strategy will only partly measure the effectiveness of the intervention. Moreover, if collected, most human capital measurements are self-reported performance measures, which makes the comparison across individuals difficult. Our work complements the literature with evidence from a vast developing country by using a data set where the results of standardized math and verbal tests are collected for all children, in and out of school.

Thirdly, we also shed light on the literature about child gender preferences. Gender biases favoring males, particularly in education, are more extensive in developing countries like China and India. Females in those countries often receive fewer investments from parents (see, for example, Oster, 2009; Jayachandran and Kuziemko, 2011; Bharadwaj and Lakdawala, 2013; Barcellos, Carvalho, and Lleras-Muney, 2014) and are likely not to reach their full potential in education, health, and personal autonomy. In our current study, the iodine policy has positive (possibly unforeseen) spillovers to females. Our study therefore also speaks to the relevance of early life conditions in explaining gender differences in socioeconomic outcomes later in life.[4]

[4]See Almond and Currie (2011) who call for work that integrates work on son preference with work on fetal origins.

Finally, our heterogeneous analysis that includes gender preferences is motivated by theoretical models of human capital formation. In such models, often dynamic complementarities between investments at different stages of childhood are considered (Cunha, Heckman, and Schennach, 2010). A handful of studies (Adhvaryu et al., 2015; Gunnsteinsson et al., 2018; Duque, Rosales, and Sanchez, 2018; Aguilar and Vicarelli, 2018; Rossin-Slater and Wüst, 2018) attempt to identify these dynamic complementarities empirically. For this, they use exogenous variation at different stages of the life cycle and generally cannot find evidence for dynamic complementarities. However, as pointed out by Malamud, Pop-Eleches, and Urquiola (2016), parents might increase investments in the child to counter the adverse effects of the initial shock. This may confound the impact of subsequent shocks. They furthermore argue that human capital outcomes for children are the result of parental preferences, the family budget constraint, and the shape of the child health production technology. This makes it challenging to interpret reduced form effect estimates. We use parental gender attitudes as a proxy for parental preferences to identify one of the channels in children's human capital formation and find that gender preferences are important in the formation of the human capital of school-aged children.

Our results point towards four observations that are relevant for the strands of literature referred to above: the relevance of parental investment responses in mitigating the effects of adverse shocks early in life; that child gender preferences are important for these investments decisions; that large scale interventions may crowd out private investments and may also affect skills dimensions initially not aimed at; large scale programs can have positive and possible unintended effects on gender equality in societies where boy preferences are important.

The rest of the paper is organized as follows. Section 2 provides a brief overview of Iodine Deficiency Disorders (IDD), the Universal Salt Iodization (USI), and related literature. Section 3 describes the data used in the analysis. Section 4 outlines empirical model and Section 5 discusses the results of the models. We zoom in on our finding of differential effects by gender in section 6. This section presents a simple model where gender preferences may differently affect parental investment in girls and boys. We introduce gender attitudes as a proxy for gender preferences and examine whether the program's effects on cognitive and non-cognitive skills vary with gender attitudes. Section 7 summarizes our findings, places these findings into context and concludes.

2 Background

Iodine is an essential component of the hormones produced by the thyroid gland and is therefore essential for human life (Zimmermann, 2011). Insufficient iodine intake causes many disorders from the fetal stage to adulthood, the most common of which is an enlargement of the thyroid gland. Although this enlargement, called goiter, is the most visible symptom of iodine deficiency, besides being inconvenient, it has no severe consequences. However, fetal exposure to iodine deficiency may lead to impaired neurodevelopment. The brain damage caused by severe iodine deficiency in this stage of life is often irreversible.

The knowledge that iodine can help prevent goiter has existed since the mid-1800's (Zimmermann, 2008). It was not until 1895 that iodine was first discovered in the thyroid gland (Baumann, 1896). Switzerland was the first country in the world to introduce iodized salt in 1922. The United States introduced iodized salt in 1924 after the executive Council of the Michigan State Medical Society officially endorsed iodized salt. In 1993, the World Health Organization (WHO) proposed a worldwide campaign to eradicate IDD. The primary intervention strategy for IDD control is Universal Salt Iodization (USI), a notably simple, universally effective, and particularly cheap instrument. The World Bank reports that it only costs approximately \$0.05 per child per year.

Historically, endemic goiter was found particularly in the mountain regions in China. For instance, in the 1940s, more than 20% of the residents of Kunming, the capital of the province of Yunnan, had goiter (Simoons, 1990). The Chinese Academy of Preventive Medicine had estimated that about 450 million people lived in iodine-deficient areas, with more than 30% of the population considered at risk of IDD (see Chen and Wu, 1998). The iodine deficiency disorders problem was acknowledged as a public health threat, and in response to this in 1993, the State Council of China announced the Universal Salt Iodization (USI) policy to eliminate IDD by 2000.

The Universal Salt Iodization was a national strategy. As the recommended strategy for controlling IDD, the USI requires that all edible salt, including salt for food processing and household use, is iodized. Accordingly, all counties throughout China should supply iodized salt (except for very few (22 out of 1800) officially approved counties). The challenge of the policy was to increase salt iodine levels sufficiently to bring the median urinary iodine concentration of children into the 100–199 µg/L range and at the same time maintaining the optimal urinary iodine concentration (MUIC) levels of pregnant women (150–249 µg/L). To reach the desired intake of iodine, the State Council enacted in October 1994 the national regulation of salt iodization. The level of salt iodization during the manufacturing process was set at 50 mg/kg in 1994.

Besides, between 1993 and 1995, a national monitoring system was built to track trends in goiter prevalence among school children aged 8-10. The monitoring was held between March and June 1995. In our empirical analyses (see section 4), we will use the outcome of this monitoring exercise as the pre-policy distribution of iodine deficiency levels across the different provinces. Note that this is a few months after the implementation of the Salt Iodization Program. Therefore, a concern may be that the cross-province variation in goiter rates does not reflect the pre-policy distribution of iodine deficiency rates. It is important to stress that the literature has documented lags of at least one year before goiter rates normalize after iodine repletion (Pardede et al., 1998; Jooste, Weight, and Lombard, 2000; Zimmermann et al., 2003).[5]

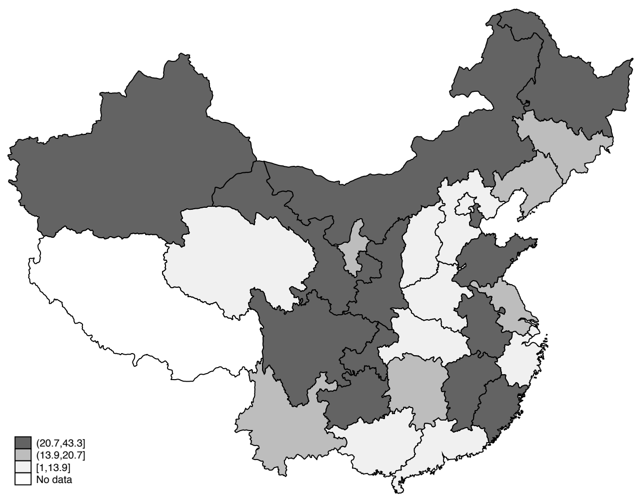

Figure 1. Goiter Distribution in 1995
Notes: Figure 1 reports goiter rates (%) (among schoolchildren aged 8-10) in 1995. Darker areas represent higher goiter rates.
Sources: National Iodine Survey 1995

Figure 1 shows the pre-policy spatial distribution of iodine deficiency levels of schoolchildren aged 8-10. Pre-policy (1995) goiter rates among children under ten years old do not differ significantly by gender (Sun, 2018). The dark areas (mostly western and northern provinces) indicate high prevalence rates (up to 43.3%), while the light areas (southeast) indicate low prevalence rates. For our empirical analyses, it is of importance to know whether the Universal Salt Iodization policy was effective in increasing the urinary iodine concentration levels in the population. We turn to this below in the data section.

[5] Pardede et al. (1998) and Jooste, Weight, and Lombard (2000) documented non-significant reductions in the size of the thyroid gland among children age 12 in South Africa one year after the introduction of iodized salt. Zimmermann et al. (2003) found in children aged 8-9 in Cote d'Ivoire even two years after the salt intervention only an eight percent reduction in the goiter rates.

3 Data

3.1 Goiter Data

The base, pre-policy, geographic distribution of goiter prevalence before the salt iodization policy (see Figure 1) came from the 1995 National Iodine Survey on goiter rates[6] among schoolchildren.[7] In each provincial survey, a multistage, probability proportional to the population size cluster sample was obtained. The county served as the primary sampling unit, and in each province, 30 counties (clusters) were selected from a county population list. In each selected county, a school was then sampled at random. Children aged 8 to 10 years at the time of the survey served as the index population. For each cluster, 40 children were selected at random from the enrollment list. All children were examined for thyroid size by palpation.[8] Therefore, the goiter rate is defined as the percentage of schoolchildren who have either Class I or Class II goiter. The sample sizes by province ranged from 1,200 to 2,400 (mean, 1,259). Our goiter data has an important advantage over goiter measures used in some recent studies like Feyrer, Politi, and Weil (2017); Adhvaryu et al. (2020) who use goiter prevalence among military recruits. This index population consists of young and healthy and as such may not be a representative measure of local iodine deficiency problems.

The survey was held every two or three years, enabling us to track the effectiveness of the Universal Salt Iodization program over time. Indeed, the program was proved to be very effective. By 2002, provinces converged to very low child goiter rates, so that provinces with high pre-eradication levels of goiter experienced the largest reductions. This is illustrated in Figure 2 that shows average goiter rates (%) across China over 1995-2002. The average goiter rate decreased from 20% in 1995 to around 5% in 2005. Figure 3 shows the post-campaign decline in goiter rate versus pre-campaign levels. Of importance for our empirical analyses is that this figure shows that the policy was effective in bringing down goiter rates for all provinces.

[6]The goiter rate is defined as the percentage of schoolchildren who have either Class I or Class II goiter. Class I goiter in normal posture of the head cannot be seen, and it is only found by palpation. Class II goiter is palpable and can be easily seen. Goiter rates in previous studies are the prevalence of Class II goiter as the information mostly comes from historical data.

[7]Although there was no representative measures of goiter prevalence before 1995, goiter prevalence of 15 areas (from 10 provinces) were measured in three consecutive years (1991, 1992, and 1993). The average goiter rate is 12.9, 15.9, and 13.6 in these 15 areas, which supports that the decline of goiter prevalence started from 1995.

[8]In some provinces, part of the children were also examined by ultrasound.

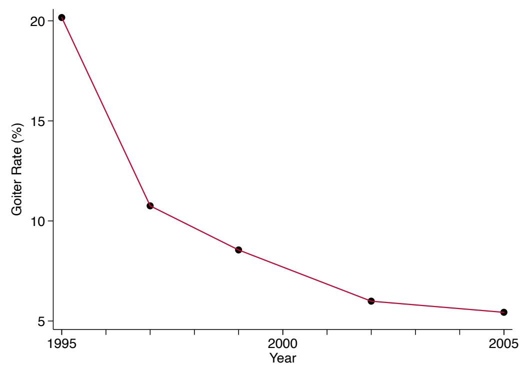

Figure 2. Goiter Rates Decline

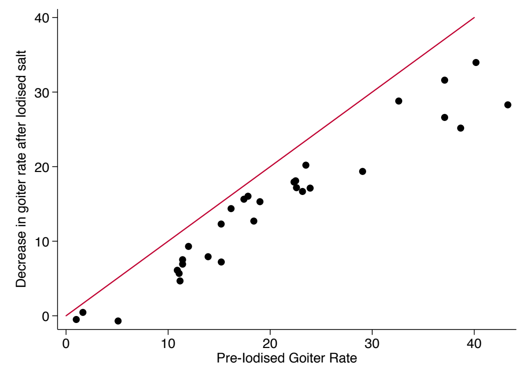

Figure 3. Before and After the Intervention

Notes: Figure 2 reports time-series data on mean goiter rate (among schoolchildren aged 8-10) across country between 1995 to 2005. Figure 3 shows the post-intervention decline in goiter rate versus pre-intervention levels across China.
Sources: National Iodine Survey 1995-2005

3.2 The Sample, Outcome Variables and Control Variables

3.2.1 The sample

The micro-level data used in this study come from the *China Family Panel Studies* (CFPS). The CFPS is a large-scale nationally representative panel survey conducted by the Social Science Survey Institute at Peking University. Three waves of the survey have been published until 2017. The CFPS baseline wave (hereafter CFPS-2010) selected a total of 14,798 households, containing 33,600 adults and 8,990 children. A second and third wave of the CFPS followed the same individual from the households in 2012 and 2014 (hereafter CFPS-2012 and CFPS-2014). A standard math and verbal tests were carried out in CFPS-2010 and CFPS-2014. For all three waves, educational attainment was recorded. Besides cognitive measures also, some non-cognitive measures were collected. We provide more information on these measures in section 6. To maintain the consistence of the sample selection, we pool the baseline wave (CFPS-2010) and the third wave (CFPS-2014) together. The data include accurate birth information, such as year, month of birth, place of birth, and whether individuals were born in a rural area. We restrict ourselves to those born in the rural areas (81% of the total population). The intervention is likely to be cleaner for those born in a rural area, as in the urban areas people had better access to micronutrient food supplements.[9] The survey also collects respondents' migration history. Migration at young ages is very low (less than 3%). The Salt Iodization policy was implemented in October 1994.

[9]The use of such supplements are also likely to vary by parental Socio-Economic Status (SES), which in itself correlates strongly with our cognitive and educational outcomes measures.

We include cohorts born between July 1990 and June 2000. At the base wave (CFPS-2010), these individuals were between 10 and 19 years old. This leaves us with 4039 children at the baseline survey. For 4021 children individuals, we have all the key information on test scores and education. The third wave (CFPS-2014) included the same math and verbal tests as the baseline wave. We, therefore, added this third wave to our baseline data. This resulted in 441 additional children. However, due to sample attrition between 2010 and 2014, we also lost 1069 individuals.[10] In the end, we have 7414 observations on schooling (3856 males and 3558 females) and 6420 observations on test scores (3310 male and 3110 female test scores). In the analyses, we link these individual observations to the pre-policy goiter rates.

3.2.2 Variables

Educational attainment is an ordered categorical variable ranging from one to four (illiterate, graduated from primary school, middle school, and high school). In the regression analysis, we use a dummy variable, indicating whether individuals have graduated from primary school. The number of years in school measures schooling. As alternative measures of human capital, we also use a math test and a verbal test designed by the CFPS team. Both tests were presented to the students, irrespective of their age. The math test was designed to test for primary and secondary school math knowledge and consists of twenty-four mathematical problems. Questions were sorted in order of increasing difficulty, and each of those questions counted for one point. Similarly, the verbal test consists of thirty-four Chinese characters based on the language textbooks again. Characters, which counts for one point each, are sorted in order of increasing difficulty. Therefore, full scores of the math and language test are 24 and 34, respectively. Note that as we measure cognition and schooling at fixed points in time, older children will perform better. It will, therefore, be important to control for age in the later regressions. Our primary empirical strategy controls for age in a very flexible way by fully exploiting the panel structure of the data set. Particularly, we always control for CFPS wave by age (dummies) interactions. Additionally, CFPS waves by cohort interactions are included to make sure we are always comparing individuals with same age. The sample also includes basic socio-demographic variables, such as age, gender, parental educations, birth order, and family size.[11] In our empirical strategy, it will be important to control for the possible mean reversion. Therefore, we supplement the set of individual controls

[10] In section 5.3, we show that the main findings are robust to the sample attrition by only using the data in the baseline.

[11] Families in the rural area were regularly exempted from the one-child policy. For example, rural married couples were allowed to have a second child if the first child was female (Zhang, 2017). The sex ratio in our sample of rural born children is about 1.07, which is lower than 1.12 (i.e., the sex ratio at birth in the 1990s documented in Jayachandran (2015)).

with a series of province-level pre-policy characteristics. All these variables are listed in Appendix A, where we also describe how variables are constructed.

Table 1 reports summary statistics of socio-demographic variables by gender for provinces with high initial (pre-policy) goiter rates (goiter rates above the median of 17%) and low initial goiter rates (goiter rate prevalence less than 17%) for cohorts born before the implementation of the program (1991-1994).[12] The Table shows that parental education is higher in regions with low initial goiter rates. The table also shows that female outcomes on test scores and education are slightly better than the outcomes for males. However, the table also shows that the differences in schooling and test scores between high and low goiter regions are larger for girls than for boys.

Table 1: Summary Statistics

	High Goiter Provinces		Low Goiter Provinces	
	Females	Males	Females	Males
Outcomes				
Educational Attainment	2.96	2.93	3.13	3.03
	[0.85]	[0.86]	[0.76]	[0.78]
Illiterate	0.053	0.066	0.016	0.0090
Primary School	0.22	0.21	0.18	0.26
Middle School	0.43	0.45	0.46	0.43
High School or above	0.29	0.27	0.34	0.31
Schooling	9.95	9.78	10.5	10.2
	[3.11]	[3.23]	[2.58]	[2.59]
Math Test Scores	15.3	15.2	16.3	15.6
	[5.67]	[5.89]	[4.90]	[5.39]
Verbal Test Scores	26.3	25.2	27.0	25.3
	[7.23]	[7.75]	[5.94]	[7.17]
Demographics				
Age	19.1	19.3	19.4	19.4
	[2.58]	[2.58]	[2.55]	[2.62]
Father's Educational Attainment	2.29	2.26	2.52	2.56
	[0.96]	[0.99]	[0.93]	[0.90]
Mother's Educational Attainment	1.70	1.72	2.10	2.04
	[0.87]	[0.87]	[0.92]	[0.90]
Birth Order	1.59	1.57	1.61	1.79
	[0.83]	[0.79]	[0.82]	[0.94]
Family Size	4.89	4.57	4.89	4.67
	[1.47]	[1.44]	[1.42]	[1.52]
Number of observations	810	895	795	926

Notes: Author's tabulations of CFPS-2010 and CFPS-2014. Sample consists individuals born in rural area between July 1990 and June 1995. We label a province as high/low goiter if its goiter rate is above/below 17% (median).

[12] In the Appendix, we provide results for the cohorts born after the implementation of the salt iodization policy (Table A2).

4 Empirical Strategy

4.1 Baseline Econometric Model

As we discussed in the previous section, salt iodization was rolled out nationwide in October 1994. Therefore, no province could serve as a pure control group. As a proxy for pre-policy iodine deficiency rates, we use province goiter rates among 8-10 years old children at the start of 1995 (see Figure 1). Like, among others, Bleakley (2010b); Adhvaryu et al. (2020), we use a difference-in-differences design. In this way, we compare trends in various outcome measures in provinces with different levels of iodine deficiency before implementing the salt iodization program.

We define someone as treated if the entire gestation period[13] is after the date of the implementation of the salt iodization program. All others are considered to be controls. In section 5, as a robustness check, we also consider alternative definitions to assign treated and controls in groups. In contrast to earlier studies that used the 1924 salt iodization policy in the U.S., the Chinese iodized salt campaign was implemented rapidly across the entire country. At the start of 1995, more than 80% of the families had already access to iodized salt. Two years later, this has increased to 95%.

During the gestational period, iodine primarily affects the fetus' neurodevelopment, with consequences for postnatal cognitive functioning (Cao et al., 1994; Zimmermann, 2011). Therefore, we focus in the first instance on the impact of the salt iodization policy on cognitive ability and school attainment of children who are affected in utero. For this, we use the following baseline regression:

$$Y_{ipt} = \beta_0 + \beta_1 Post_t \times Goiter_p + X_{ipt}\rho + \delta_p + \gamma_t + \epsilon_{ipt} \tag{1}$$

where outcome Y_{ipt} is either the logarithm of the cognitive test scores (math and verbal), educational attainment as well as the number of schooling years for individual i, who was born in province p, in year t. $Post_t$ indicates whether the individual was conceived after the introduction of iodized salt. $Goiter_p$ is a measure of pre-eradication endemicity in individual i's province of birth. As mentioned above, we use goiter rates collected in the National Iodine Survey held at the start of 1995. The mean goiter rate is 20%, and the standard deviation is 12%.

The vector X_{ipt} includes individual characteristics: parents education and family size, and mean-reversion controls.[14] First, we follow Bleakley (2010a) and construct the mean-

[13] As is generally done in the literature, we use a nine-month gestation period.

[14] If the oldest cohorts had high Iodine Deficiency Disorders and low human capital because of some mean-

reversion control by interacting provincial average educational attainment in the 1990 Census with the dummy variable $Post_t$. Second, we control for many baseline province pre-treatment characteristics interacted with cohort dummies. In our baseline model, we control for hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990, again all interacted with cohort dummies. In a robustness check (section 5.3), we additionally control for the number of schools per capita in 1991, poverty rates in 1993, and average household income in 1991 all interacted with cohort dummies.[15]

To take into account age effects and the year of survey effects, CFPS wave by age (dummies) interactions are also included. Additionally, CFPS waves by cohort interactions are included to account for differential trends in the outcome variables. δ_p and γ_t are province and birth cohort fixed effects. δ_p and γ_t are province and birth cohort fixed effects. Note that the birth cohort fixed effects are important here as our outcome variables are measured in 2010 and 2014, which implies that those exposed to the salt iodization are much younger than the controls. We also control for region-specific linear trends in all models.[16] We run specification (1) separately for males and females. We also consider alternative specifications in section 5.3. Of prime interest is the continuous treatment variable $Post_t \times Goiter_p$ that proxies potential iodine exposure. Recall from Figure 3 that the salt iodization policy was very effective in reducing goiter rates in all provinces to very low levels. So while the parameter β_1 in a strict sense is the intention-to-treat effect, the high compliance rates make it very close to the treatment effect.

4.2 Dynamic Specification

Since our estimates use the cross-province convergence in goiter rates created by the introduction of iodized salt (Figure 2 and 3), convergent pre-trends across high and low-base goiter rate provinces prior to 1995 are a concern. Therefore, we also use an event study design to test for the common pre-trends assumption formally. More specifically, we run the following regression:

$$Y_{ipt} = \beta_0 + \sum_{t=1990}^{2000} \beta_t \times Goiter_p + X_{ipt}\rho + \delta_p + \gamma_t + \epsilon_{ipt}, \tag{2}$$

reverting shock. We might expect human capital gains for the subsequent cohorts, even in the absence of a direct effect of the salt iodization policy on productivity.

[15] The provincial poverty rates in 1993 are obtained from Woo et al. (2004). The average household income in 1991 is constructed from the National Fixed Point Survey.

[16] The regions consist of several provinces. See Table A1 of the Appendix A for the precise definition.

where β_t gives the cohort-specific relationship between pre-eradication endemicity and later-childhood outcomes.[17] If salt iodization affected the human capital formation of exposed cohorts, these effects should be visible in a break from pre-existing trends in β_t. This method will also shed light on the partial effects of iodine exposure in late childhood (rather than in utero) if such effects exist.[18] Note that all individuals born in 1995 or later are exposed to iodized salt from conception onward. Individuals born in 1994 experience higher iodine intake in their year of birth; thus, this cohort is partially exposed to higher iodine levels in utero and fully exposed from birth onward. Individuals born in 1993 experience higher iodine from age one onward, and those born in 1992 experience higher iodine from age two onward. Since we normalize the 1994 cohort coefficient to zero, our analysis essentially tests for differential effects of exposure relative to exposure at age one and older. If there are additional benefits to having access to iodine between conception and age 1, we would expect the coefficients β_{1995+} to be positive. Similarly, if iodine at age 1 has an additional benefit relative to iodine exposure at age two or older, we would expect coefficients β_{1993-} to be negative.

5 Results

5.1 Baseline Results

Table 2 reports the main results of two separate regressions of our basic model (Equation 1): one for men in Panel A and the other for women in Panel B. In all the regressions discussed in this section, the coefficients of interest are the post-by-goiter rate interaction, which represent the effect of salt iodization on our outcomes of interest. The coefficients in all tables have been multiplied by 12, the inter-quartile (25-75) range of the goiter distribution. The size of the effect is scaled to be the effect of moving from a relatively high goiter province to a low goiter province. Although the following tables only report the coefficient of interest, in all specifications, we include controls for province and year of birth fixed effects, birth order, family size, parents' characteristics, region-specific linear trends, a series of mean reversion controls, age and the survey year controls. Standard errors are clustered at the province-of-birth level to allow for arbitrary correlation of the errors for individuals

[17] In practice, β_{1994} represents estimates of individuals born between July 1994 and June 1995, and β_{1995} for the cohort born between July 1995 and June 1996, etc. We made such adjustments because individuals born after July 1995 were conceived after implementing the salt iodization policy. Similarly, the 1993 cohort are those born between July 1993 and June 1994.

[18] See Zimmermann (2011) for a comprehensive summary of the role of iodine in human growth at different stages in life. For example, neonatal iodine deficiency may cause endemic cretinism. Deficiency during childhood and adolescence may impair mental functioning and delay physical development.

born in the same province. We also report two-way clustered standard errors by province and family (in parenthesis) to take account of situations where there are multiple children in a family. The standard errors of two-way clustering are almost identical to the usual cluster robust standard errors. Given that our data only contain 28 provinces, we produce statistical inference based on the wild-bootstrap approach (Cameron, Gelbach, and Miller, 2008). The associated wild-bootstrap standard errors (in angle brackets) turn out to be similar to the usual cluster robust standard errors. For all four human capital measures, we identify significant effects of the intervention for females. A one standard deviation decrease in the pre-intervention goiter rate is associated with about 4% increase in the probability of graduating from primary school, a 10% increase in schooling, 14% increase in math and 12% increase in verbal test scores. For males, the coefficients are substantially smaller in magnitude and not significant for any of the outcome variables.

Table 2: Iodine Exposure and Human Capital Attainment

	(1) Math Test ln(scores)	(2) Verbal Test ln(scores)	(3) Primary School	(4) Schooling ln(years)
Panel A: Males				
Post × Goiter	0.0298	0.0605	0.0171	0.0228
	[0.0508]	[0.0511]	[0.0226]	[0.0411]
	(0.0566)	(0.0666)	(0.0234)	(0.0370)
	⟨0.0553⟩	⟨0.0492⟩	⟨0.0210⟩	⟨0.0428⟩
Mean of Dep. Var.	2.517	3.083	0.497	2.108
Observations	3310	3310	3791	3654
Panel B: Females				
Post × Goiter	0.137	0.117	0.0382	0.108
	[0.0372]***	[0.0556]**	[0.00811]***	[0.0202]***
	(0.0310)***	(0.0468)**	(0.0134)***	(0.0185)***
	⟨0.0312⟩***	⟨0.0507⟩**	⟨0.0122⟩***	⟨0.0152⟩***
Mean of Dep. Var.	2.543	3.170	0.528	2.123
Observations	3110	3110	3510	3386

Notes: Each coefficient is from a separate regression. All regressions control for fixed effects specific to birth province and birth year, birth order, family size, parents' education, region-specific linear trends, survey wave by age interactions and survey wave by cohort interactions. Mean-reversion controls include provincial average educational attainment in the 1990 Census interacted with the dummy for treated cohorts, hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. Two-way clustered standard errors by province and family appear in parenthesis. P-values based on wild-bootstrap approach (Cameron, Gelbach, and Miller, 2008) with 999 replications appear in angle brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.

5.2 Results from the Event Study

Crucial for the causal interpretation of our coefficients is the assumption of common pre-intervention province trends in the outcome variables. In order to test for differential pre-trends, we run regressions of the form specified in Equation (2). Figures 4, 5, 6 and 7 plot the estimated coefficients (β_t) of the event study for females. Results for males are reported in Appendix B (Figure B1, B2, B3 and B4). The figures show that the trends leading up to the year of the intervention are identical and insignificant from the 1994 effect. This gives confidence in the validity of the common trend assumption. Figures 4 and 5 show that math and verbal scores of females increase shortly after the intervention with about 10%. For educational outcomes (Figure 6 and 7), we see a similar picture (about 10% for both educational attainment and years of schooling, respectively). While there is some variation in the effect size of the program across the cohorts, the 95 percent confidence intervals overlap for all cohorts. This suggests that females born in the later cohorts obtain comparable benefits from the adoption of iodized salt. The results for males in Appendix B point at insignificant effects for all outcome variables and all cohorts.

5.3 Robustness Analysis

There are several reasons why trends in educational outcomes across birth cohorts might differ across provinces. In this section, we explore some additional specification checks to make sure that our baseline results can be interpreted as causal. Table 3 presents these additional specifications/checks. Here we restrict ourselves to the results of robustness checks for females. The same robustness checks for males are reported in Table B1 of Appendix B.

To ensure that the logarithm transformation does not drive our finding, we run the same regression using raw math and verbal scores as dependent variables. The results reported in Panel A show that the conclusions from the baseline models remain to hold and that the functional specification does not drive the findings.

Another concern of our identification strategy is mean reversion. In the baseline, we control for many baseline province pre-treatment characteristics interacted with cohort dummies. To further decrease the concern for mean reversion, we also control the number of schools per capita in 1991, poverty rates in 1993, and average household income in 1991 all interacted with cohort dummies. The provincial poverty rates in 1993 are obtained from Woo et al. (2004). The average household income in 1991 is constructed from the National Fixed Point Survey. Instead of region-specific linear trends, we control for the region of birth by birth year interactions. The results of this exercise, reported in Panel B, show that the effects remain significant and are quantitatively similar (if not, larger) than the baseline

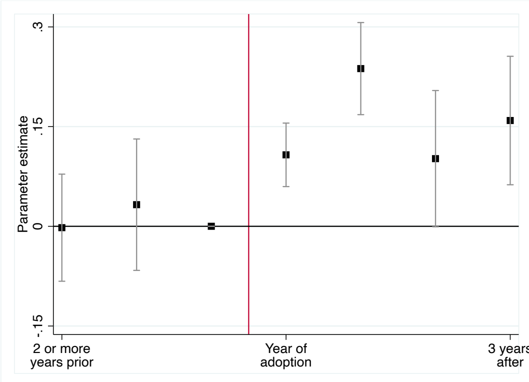

Figure 4. Math Test Scores

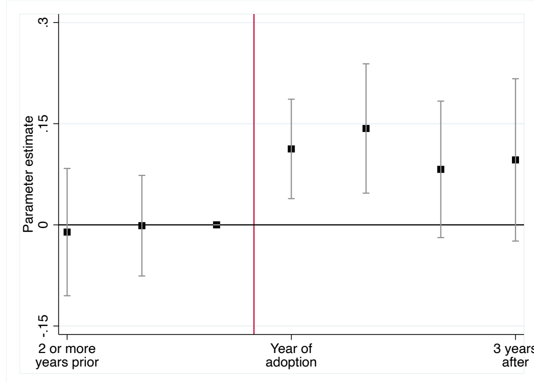

Figure 5. Verbal Test Scores

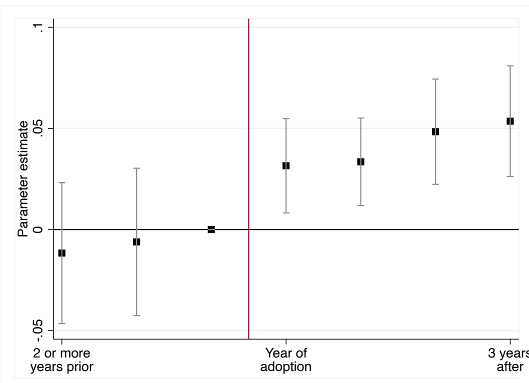

Figure 6. Primary School

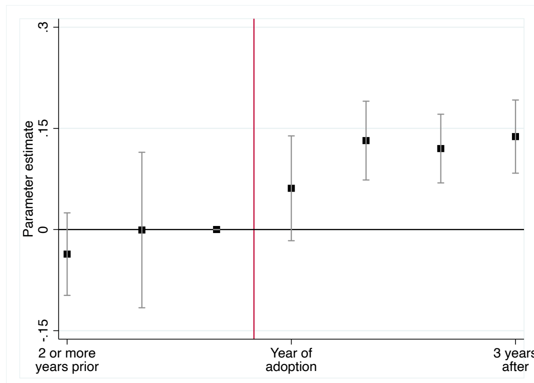

Figure 7. Schooling

Notes: The sample includes all female respondents from two waves of the survey (CFPS-2010 and CFPS-2014). Each point reflects the coefficient estimated on an interaction term between the birth year (compared to 1995) and the pre-intervention (base) level of the goiter rate in the birth-province. Capped spikes represent 95 percent confidence intervals. All models condition upon birth province and birth year fixed effects and the full set of controls used in our main analysis.

estimates. A one standard deviation decrease in the pre-intervention goiter rate is associated with a roughly 15% increase in cognitive ability and 11-13% in educational outcomes for females.

In section 3, we assigned individuals into the treatment group if the full gestation period was after October 1994. All others, the partially exposed and those who were not exposed, were placed in the control group. This assumes that the gestational period was precisely nine months and that those whose gestational period partly lies after October 1994 are not exposed in utero. This classification is perhaps a bit too conservative. Panel C reports the results using the sub-sample where the partially exposed children (born in April, May, and June 1995) are dropped from the analyses. The results in Panel C show that this hardly affects the estimates.

The verbal and math test remained the same over the 2010 and 2014 wave. In the basic

Table 3: Robustness Checks (Female)

	(1) Math Test ln(scores)	(2) Verbal Test ln(scores)	(3) Primary School	(4) Schooling ln(years)
Panel A: raw test scores				
Post × Goiter	0.814	0.667	0.0399	0.493
	[0.239]***	[0.439]	[0.00783]***	[0.108]***
Mean of Dep. Var.	13.90	25.23	0.530	8.730
Observations	3110	3110	3522	3386
Panel B: additional controls				
Post × Goiter	0.108	0.106	0.0230	0.0842
	[0.0437]**	[0.0509]**	[0.0224]	[0.0346]**
Mean of Dep. Var.	2.543	3.170	0.530	2.123
Observations	3110	3110	3522	3386
Panel C: drop partial exposed group				
Post × Goiter	0.124	0.113	0.0386	0.106
	[0.0395]***	[0.0536]**	[0.00754]***	[0.0183]***
Mean of Dep. Var.	2.540	3.169	0.529	2.120
Observations	3026	3026	3426	3297
Panel D: only using baseline wave 2010				
Post × Goiter	0.129	0.105	0.0442	0.100
	[0.0310]***	[0.0468]**	[0.0142]***	[0.0185]***
Mean of Dep. Var.	2.491	3.128	0.343	1.975
Observations	1861	1861	1906	1906
Panel E: small sample window				
Post × Goiter	0.137	0.136	0.0322	0.0899
	[0.0278]***	[0.0495]**	[0.0106]***	[0.0201]***
Mean of Dep. Var.	2.580	3.196	0.553	2.146
Observations	2529	2529	2877	2755

Notes: Each coefficient is from a separate regression. All regressions except Panel B use the same controls as the baseline model in Table 2. In Panel B, we control for birth-region and birth-year specific interaction instead, and we additionally control for schools per capital in 1991, poverty rates in 1993, and average household income in 1991 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2010 and CFPS-2014

specification, we pooled the observations from the 2010 and the 2014 wave. Individuals in our analysis are in their late teen ages/early-adulthood, and therefore as we do in the regressions controlling for age is important. However, using information from all waves has a threat that individuals might learn from the test results in the first wave. This makes the second wave test scores inappropriate as a measurement of development in cognition. As a check, we run regressions using only the data from the first wave of the CFPS. Results of using only first wave in 2010 is also informative on the impact of sample attrition between the first wave in 2010 and the last wave in 2014. Indeed, we lost around 30% percent of observations due to attrition (see section 3). The results in Panel D show that the results remain close to our baseline specification and confirm our main findings.

We also used a slightly smaller sample window by restricting individuals born between July 1991 and June 1999. The advantage is a lower attrition rate in wave 2014 as older cohorts (born in 1990) are more likely to have left home. Moreover, the trimming of the left and right tail of the age distribution makes the sample more homogeneous. A downside of this age/cohort restriction is a smaller sample size and hence less power. Still, however, all four estimates in Panel E remain to be significant and close to the baseline estimates in Table 2.

We also performed a placebo test, where use parental education (which was used as a control variable in the main specifications) as outcome variables. The idea of this test is that parental education should not be affected by future exposure to iodized salt, and if we do find an effect, it may suggest that our results may be driven by parental background. Indeed, both mother and father's education (measured by education attainment or years of schooling) are not affected by future exposure (results available upon request).

Finally, we further verify our identification assumptions and demonstrate the statistical power of our inferences by conducting falsification tests where we assign a pseudo-treatment. More precisely, we randomly assign province of birth and thus pre-intervention goiter rates to each respondent in our sample. If our identification strategy is valid, we would expect estimates using those pseudo-samples to be centered around zero. We can then confront our baseline estimates with the results from the pseudo-sample. In Figures 8, 9, 10 and 11, we plot the distribution of the t-statistics from 5,000 estimated pseudo-treatment effects on educational attainments, schooling, math and verbal ability, respectively. As expected, all four distributions are centered around zero. Together, these results imply that assumptions in our empirical model are unlikely to be violated. To address the statistical power of our model, we mark within the distribution of pseudo-treatment effects the location of the t-statistic of the baseline treatment effects in Table 2. We also report at the top of each figure the share of the pseudo-treatment t-statistics that exceed the actual t-statistic of the baseline model (in absolute values).[19] These p-values give confidence in our design and statistical power of our exercises.

[19] These p-values can be seen as alternatives to the p-values obtained from our clustered standard errors reported in Table 2.

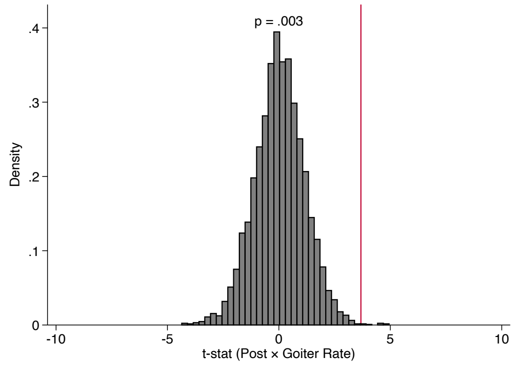

Figure 8. Math Test Scores

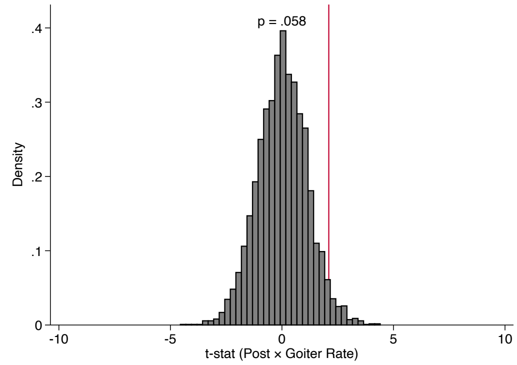

Figure 9. Verbal Test Scores

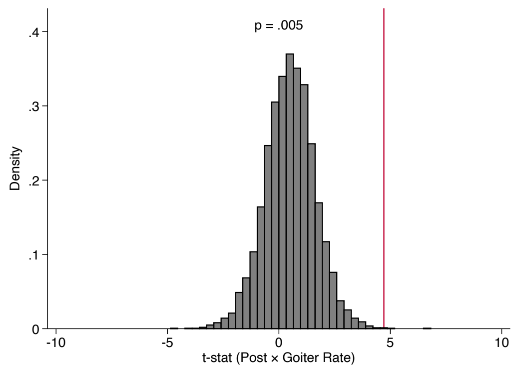

Figure 10. Primary School

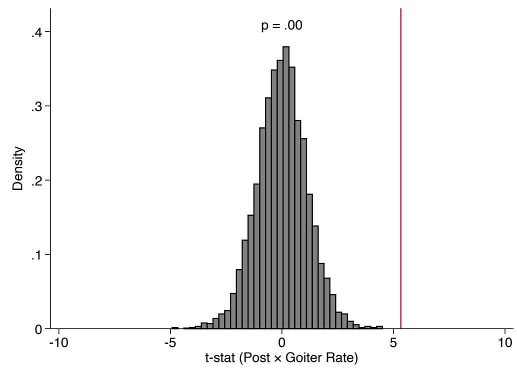

Figure 11. Schooling

Notes: Pseudo-treatment vs. actual policy intervention: the distribution of t-statistics resulting from 5,000 random assignments of treatment to individuals, as well as the t-statistics from the actual treatment through the policy intervention (red line). "p-values" report the share of the pseudo-treatment t-statistics that is larger than the actual t-statistics.

5.4 Endogenous Sex Selection

In this section, we address the potential threat posed by endogenous sex selection. Goiter rates may influence the gender ratio for biological reasons. Alternatively, there could be parental motives related to goiter prevalence rates to sex select. In the presence of ultrasound techniques, such selection effects are not inconceivable. In either case, this would imply that the marginal girl born in a high goiter area would be different from the marginal girl born in a low goiter area, which may pose a threat to the interpretation of our findings.

To rule this out, we test directly whether the provincial sex ratio is correlated with iodine deficiency (goiter prevalence rates). More importantly, we test whether the association changes after 1994. Following Edlund et al. (2013) we use the Census of 2000 to calculate the sex ratio by the birth year (1987 to 1999) at the province level. Linking this data with provincial goiter rates, we display in Figure 12 the results of an event study regression that relates provincial sex ratios to goiter prevalence rates. As may be clear from the graph, there is no association between the sex ratio and iodine deficiency, and the association does not change after introducing the salt iodization policy.

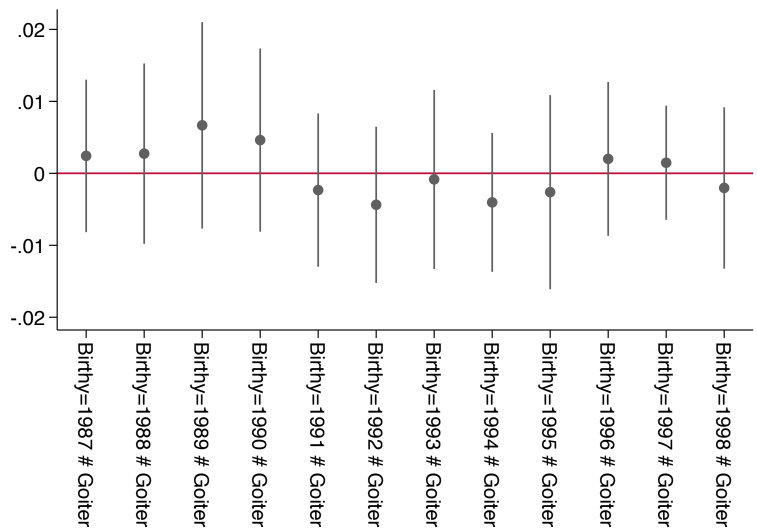

Figure 12. Sex Ratio and Iodine Deficiency

Notes: Sex ratios are aggregated from microdata of the 2000 Census. Following Edlund et al. (2013), we calculate the sex ratio at province level for each birth year between 1987 and 1999. In the event study regression, we control for province fixed effects and year of birth fixed effect. Therefore, cohort who were born 1999 is the baseline group and the coefficient of the cohort 1999 is then omitted in the graph. Standard errors are clustered by province.

5.5 Comparison with Other Cohort-Based Iodine Studies

Some recent studies (Field, Robles, and Torero, 2009; Feyrer, Politi, and Weil, 2017; Adhvaryu et al., 2020) also analyze the long-run impacts of iodine deficiency in early-life. Field, Robles, and Torero (2009) was one of the first to provide evidence at the micro-level of the effect of iodine availability in utero on schooling attainment in Tanzania. Feyrer, Politi, and Weil (2017); Adhvaryu et al. (2020) exploit a nationwide salt-iodization program initiated by the public health authorities in Michigan in 1924.[20] Similar to our study, they use pre-program geographical information on goiter prevalence along with the time variation in the introduction of iodized salt to assess the causal effect of iodine on later life outcomes. Feyrer, Politi, and Weil (2017) find strong effects for males: iodized salt in utero leads to a 15 point increase in I.Q. Interestingly, Adhvaryu et al. (2020), using census data, find no effects for males, but strong increases in income (11%), labor force participation (0.68%) and

[20]Morton's salt, the largest producer in the U.S. at that time, began selling iodized salt in the fall of 1924.

full-time work (0.9%) for females. Field, Robles, and Torero (2009) also find stronger effects for girls. Like Adhvaryu et al. (2020), we only find effects for females. Relating our findings to Field, Robles, and Torero (2009) and Adhvaryu et al. (2020), we find that the introduction of iodized salt leads to 0.441 additional years of schooling (see panel A of Table 3). This estimate is in line with Field, Robles, and Torero (2009), who find that the iodine supplement program increased schooling years with 0.35-0.56 years. Wang (2013) finds for China that one year of additional schooling raises income with 15%, which translates to an about 5% (0.441×0.11) increase in income for females in our study.

Our strong effects of iodine in utero for females is consistent with the medical literature that posits that female fetuses are more sensitive to maternal thyroid deficiency than male fetuses Zimmermann (2011). However, this does not imply the absence of effects for males (see, for instance, the findings of Feyrer, Politi, and Weil (2017) and Field, Robles, and Torero (2009)). In developing countries, gender differences in socio-economic outcomes are bigger than in developed countries, and cultural gender norms may contribute to gender differences in human capital outcomes. These gender norms may already have an impact on children very early in life and might explain heterogeneous gender effects found in the literature on long-run effects of early life shocks.[21] Indeed, parents' preferences for boys make these parents to desire more sons than daughters, but it may also result in parents choosing to invest more in sons than in daughters (Jayachandran, 2015). The former mechanisms may result in a male-skewed sex ratio. Almond and Edlund (2008); Abrevaya (2009) find that son preferences in Asian immigrants to the U.S. seem to persist with changing the economic environment, suggesting a strong role for culture (i.e., preferences). Important for our study is that son preferences may imply not only more material and non-material resources in boys post-birth but also that parental investments post-birth may mitigate some of the adverse effects of iodine deficiency in boys more than in girls. This would be consistent with our finding of the strong effects of the salt iodization policy for girls and no effect on boys. Below we discuss the role of parental investments, which helps to explain the mixed finding in the literature concerning gender.

[21] This was recently also argued by Dinkelman (2017), who refers to the effects of local shocks to the environment in the Asian context that affect resource availability. Findings for Indonesia (Maccini and Yang, 2009) and India (Pathania, 2007) point at more negative effects for girls.

6 More on Gender Differences

6.1 Conceptual Framework

To explicitly consider possible interactions between gender preferences in receiving parental investment and endowment early in life, we consider a slightly adjusted version of the model of child human capital formation (Yi et al., 2015; Almond, Currie, and Duque, 2018). In the model, each child has two components of human capital: cognitive skills (θ^C) and non-cognitive skills (θ^N). Parents can influence the formation of human capital by making investments in the child's skill dimensions I^k, $k \in \{C, N\}$. The optimal investment decision follows from a maximization of the parents' utility function:

$$U = U(c, q) = (1 - \alpha_s) \log c + \alpha_s \log q(\theta^N, \theta^C), \tag{3}$$

subject to a budget constraint and the child's human capital production function, where $s \in \{boys, girls\}$, c parental consumption and q child quality. The parameter α_s captures the preferences for child quality and may vary across families. Son preferences imply higher values of α_s for boys, $\alpha_{boys} > \alpha_{girls}$. From the optimization of the model, it follows that the optimal investment depends on the child's initial endowments, the price of consumption, the price of investments in skills and preferences α_s. Note that investments could be broadly interpreted. It could be well-targeted efforts aimed at improving specific skills as well as the provision of an environment (e.g., food, housing, attention, etc.) that foster child well-being and child outcomes. In this multidimensional model of human capital formation, investment strategies of parents depend not only on parental preferences but also on the production technology and the budget constraint.[22]

Following Yi et al. (2015); Almond, Currie, and Duque (2018), the total effect of a shock early in life (e) on cognition can be decomposed into two parts:

$$\underbrace{\frac{\mathrm{d}\theta^C}{\mathrm{d}e}}_{A} = \underbrace{\frac{\partial \theta^C}{\partial e}}_{B} + \underbrace{\frac{\partial \theta^C}{\partial I^C}}_{C} \times \underbrace{\frac{\partial I^C}{\partial e}}_{D}. \tag{4}$$

A similar expression can be obtained for non-cognitive outcomes (θ^N). The term (A) on the left-hand side of (4) is the total effect of an early-life shock and corresponds to the usual reduced form estimates in the empirical literature. In our study, before the salt-iodization policy, the term (A) would be the effect of iodine deficiency, i.e., the pre-intervention cognitive

[22] Almond and Mazumder (2013) give examples that show that compensation for an early-life shock is optimal in some settings and not in others. Moreover, the absence of behavioral responses could also be the result of a family being financially constrained. See Almond and Mazumder (2013) for more details.

outcomes in late childhood due to geographical variation in iodine levels. The first term on the right-hand side (B) is the biological effect that directly operates through the production function. The second term ($C \times D$) is a behavioral effect from the parental investment response, where (C) is the productivity effect of the investment (the marginal Efficiency of Investment) and (D) the resource allocation effect. The resource allocation effect (D) depends on parental preferences α_s. The total effect of the shock (A) will be less than the biological effect (B) in case parents respond with investments to counter the adversity of the early-life shock. High values α_s imply more weight to child outcomes and imply higher levels of parental investment to counter the early-life shock. Conversely, given a biological effect (B), a higher weight (α_s) for boys than for girls implies more investments for boys than for girls and hence smaller reduced form effects of iodine deficiency for boys at later ages. This is in line with the findings of some empirical studies.

The salt iodization policy neutralizes geographically determined iodine deficiency rates. The policy has, therefore, two consequences. First, it reduces the negative cognition effects of iodine deficiency for both genders. In the presence of gender preferences, the reduction in female disadvantage in cognition due to the reform will be more substantial than for boys (holding the biological effect (B) constant). This is consistent with our finding of positive and significant cognition effects of the salt iodization policy for girls and small and insignificant effects for boys. Second, the salt ionization program will alter parental investment decisions. The need to invest in cognition is reduced or may even be absent, and parents may divert (part of the) investments into other dimensions of human capital (notably non-cognitive skills). For example, parents may set a particular target for schooling for their children. When after the salt iodization program targets for schooling are met, parents may shift some of their investment skill dimensions.

Although the literature on non-cognitive skills in developing countries is still very limited, there are several reasons to hypothesize that parents will invest in non-cognitive skills. Firstly, parents can easily observe emotions, mental and physical health. Secondly, parents might believe that non-cognitive skills could improve the child's welfare. Although there is little evidence about the returns to income from non-cognitive skills in a developing country, parents may be convinced that non-cognitive skills can benefit their child's welfare via other dimensions (e.g., the marriage market).

Although the literature on non-cognitive skills in developing countries is still limited, there are several reasons to hypothesize that the parental can and will respond to and invest in non-cognitive skills. Firstly, non-cognitive skills might be quite salient. Parents can easily observe their child's mental health and emotion, mainly targeted by the CES-D questionnaire. Secondly, parents might reasonably believe that non-cognitive skills could substantially affect children's welfare. Even there is very little evidence around the returns to non-cognitive skills in a developing country context in terms of income, parents can still reasonably conclude that non-cognitive skills can benefit their children in other dimensions (e.g., the marriage market).

Note that this does not imply that cognitive and non-cognitive skills are substitutes and that this behavior does not contradict skill complementarity in the sense of Heckman (2007).[23] Indeed, a prediction from the model (Yi et al., 2015; Almond, Currie, and Duque, 2018) is that parents can compensate and reinforce initial shocks along different dimensions of human capital. Or, related, via "cross-productivity" (Cunha et al., 2006), changes in one dimension of human capital may also affect the accumulation of other dimensions.[24] Therefore, although the medical literature primarily documents neuro-developmental impairments as a consequence of in-utero exposure to iodine deficiency (Zimmermann, 2011), it cannot be ruled out that elimination of iodine deficiency also impacts non-cognitive dimensions of human capital (such as physical health, personality traits, risk attitude, etc.). The effect (sign and magnitude) of the salt iodization on non-cognitive outcomes is ultimately an empirical matter.

We use the baseline specification (1) to estimate the effect of the salt iodization policy on five different measures of non-cognitive skills. The first four measures are obtained from the 20 questions in the CES-D module of CFPS-2012. See Appendix C for details about the construction of the four measures. The fifth measure of non-cognitive skills is derived from a question in CPFS-2014: "How well are you getting along with others?". The results of the difference-in-differences regressions for the non-cognitive skill measures are displayed in Table 4. The table shows that the salt-iodization policy significantly improved retarded somatic activity, depressed affect, interpersonal problems, and social skills for boys. The corresponding coefficients for females are much smaller and insignificant. This finding is consistent with the idea that when the salt-iodization policy eliminates the cognitive disadvantage at birth, parental investments in boys that were initially geared towards cognitive

[23]It needs to be pointed out that here that dynamic complementarities, as discussed in the literature by Heckman (2007), only imposes restrictions on the functional form of the human capital production technology. Importantly, the dynamic process of human capital formation is jointly determined by parental preferences, budget constraints, and the production technology. So, even if there are dynamic complementarities in the technology of human capital formation such that "capabilities beget capabilities", parents can still respond endogenously by mitigating or reinforcing disadvantages early in life. Similar arguments have been made by Malamud, Pop-Eleches, and Urquiola (2016) and were confirmed by a series of empirical studies (Adhvaryu et al., 2015; Gunnsteinsson et al., 2018; Duque, Rosales, and Sanchez, 2018; Aguilar and Vicarelli, 2018; Rossin-Slater and Wüst, 2018).

[24]Note, however, that in our simple model there is no role for "cross-productivity" and this is thus not captured in the decomposition (4). An example of decomposition which includes "cross-productivity" can be found in Grönqvist et al. (2018).

skills are now diverted to other dimensions of human capital. Two alternative explanations are biological pathways and "self-productivity of skills" denoted by Cunha, Heckman, and Schennach (2010). To the best of our knowledge, there is no medical evidence that iodine deficiency in utero is related to the non-cognitive skills development. Therefore, the biological pathway is not likely to be the main explanation. And recall that we do not find significant impacts of the policy on boy's cognitive abilities, "self-productivity of skills" is also not likely to explain the positive effects on the non-cognitive skills of boys. However, it cannot be ruled out that primarily, the production technology of cognitive and non-cognitive skills and their interaction drives our findings (Almond, Currie, and Duque, 2018). We also used the total CES-D score as a measure for non-cognitive skills. We find a sizable and significant coefficient of -0.912 (with a p-value of 0.044).

Table 4: Iodine Exposure and Non-Cognitive Skills

	(1) Somatic Complaints	(2) Depressed Affect	(3) Positive Affect	(4) Interpersonal Problems	(5) Social Skills
Panel A: Males					
Post × Goiter	-0.354	-0.453	0.0612	-0.147	0.772
	[0.134]**	[0.162]**	[0.205]	[0.0765]*	[0.447]*
Mean of Dep. Var.	3.331	2.542	4.783	0.555	22.33
Observations	1216	1218	1218	1218	950
Panel B: Females					
Post × Goiter	0.104	-0.0147	0.111	-0.0689	-0.0150
	[0.195]	[0.233]	[0.151]	[0.0702]	[0.474]
Mean of Dep. Var.	3.539	3.259	4.898	0.657	22.46
Observations	1180	1180	1179	1180	915

Notes: Each coefficient is from a separate regression. All regressions control for fixed effects specific to birth province and birth year, birth order, family size, parents' education, region-specific linear trends and age. Mean-reversion controls include hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2012 and CFPS-2014.

6.2 Gender Attitudes

Although differences in parental preferences can rationalize gender differences in the effects of the iodization policy, we cannot rule out the possibility that biological differences between men and women entirely drive all such gender differences. Therefore, in this section, we put the focus on within-gender analyses. We exploit plausibly exogenous variation on the resource allocation effect (D) generated by parental preferences (α_s in the utility function). Assuming that for a given gender, the biological effect (B) is independent of parental preferences (α_s), then the total reduced form effect (A) will not vary with parental preferences if there are no behavioral effects. On the other hand, reduced form estimates that vary with gender preferences hint at the relevance of behavioral responses of parents.

The assumption that gender preferences play no role in the biological effect requires more discussion. With the introduction of ultrasound techniques, parents could respond prenatally, either with selective abortion or with increased antenatal investments when a male fetus is identified (see Bharadwaj and Lakdawala, 2013). In the latter case, we take the position that antenatal investments are included in parental investment decisions. As regards selective abortions, our empirical analyzes in the next section rely on within gender variation in parental preferences. This allows us only to need the much weaker assumption that for a given gender, the biological effect does not depend on α_s. We provide a test based on sex-ratios that supports this assumption later in section 6.3.

We do not directly observe gender preferences and rely on a plausibly exogenous proxy for parental gender preferences. We take the advantage that the CFPS-2014 includes a module with several questions on gender equity attitudes.[25] This module covers topics on gender roles within the household and in public life and asked respondents whether they agree with six statements phrased against gender equality and women empowerment. The response categories ranged from 1 "Strongly disagree" to 5 "Strongly agree". See Table C1 of the Appendix C for the statements and average responses for adult males and females. The gender attitude index at the two bottom rows are the mean and normalized mean of the individual responses for the six statements. A lower gender index means more gender-equitable views. Ideally, we would like to have gender attitudes measured 10 to 20 years before so that they coincide with the period during which investments were made. Unfortunately, this information is not available. Therefore, we approximate gender attitude measured in 2014, by which we have to assume that gender attitudes differences are relatively stable over time. For example, Attané (2012) found that the percentage of women(men) who agree with "Men are turned toward society, women devote themselves to their family" actually increased by only 4.4% (7.7%) from 50.4% (53.9%) from 2000 to 2010. The fact that percentage remains stable both in men and women reveals the deep-seated internalization of gender inequality.

Gender attitudes have drawn great interest in the recent literature (see Dahl, Kotsadam,

[25] An alternative proxy could be the sex ratio (see Edlund et al., 2013, for instance). Selective abortion may be higher in regions where preferences for boys are stronger than in other regions. We do not use sex ratio as a proxy for gender preferences as it is not a measure of son preference *per se*, but rather the realization of the family's son preference combined with the preferences over the family size Jayachandran (2015).

and Rooth, 2017; Dhar, Jain, and Jayachandran, 2018; Dossi et al., 2019). For instance, Dossi et al. (2019) explicitly link son preference to under-performance of girls in the U.S. We follow the design of Dossi et al. (2019) to examine the association between gender attitudes and parental investment. The parental investment is measured by an unweighted average of the z-score of 5 variables: breastfeeding practices, whether the child went to kindergarten, how often the parent read to their child, how often the parent buys books for their child; and how often they travel with their child. Unfortunately, who do not observe such parental investment score for the parents of our target sample of children born between 1991 and 2000. We do, however, observe the parental investment score for children younger than 6 in CPFS.

We regress the parental investment score on gender attitude, an interaction of gender attitude with gender, and a set of individual characteristics (Table C4 in Appendix C). We find gender attitudes to be strongly associated with parental investments in children, and boys' investments are higher in areas where preferences favoring boys are stronger. This finding remains robust when we exploit within province variation in the gender attitude variable by adding province fixed effects. This is important for the next sub-section, where we exploit within province variation in gender attitudes to identify the effect of the iodization policy on cognitive and non-cognitive outcomes by gender.

6.3 Gender Attitudes and the Effect of the Salt Iodization

Assuming that the biological effect does not depend on gender preferences α_s, the model in section 6.1 hypothesizes that systematic variation of the reduced form estimate (A) with gender preferences hints at an important role for parental gender preferences in human capital investments in children. Gender preferences can be incorporated by gender specific preference weights α_s and to test for this we specify the following triple-difference equation that we separately estimate per gender:

$$Y_{ijpt} = \beta_0 + \beta_1 Post_t \times Goiter_p \times GA_j + \beta_2 Post_t \times Goiter_p + \beta_3 Post_t \times GA_j + \beta_4 GA_j + X_{ipt}\rho + \delta_p + \gamma_t + \epsilon_{ipt} \quad (5)$$

Y_{ijpt} is the human capital outcome of child i, living in village/community j of province p at time t. $Post_t$, $Goiter_p$ and X_{ipt} are defined as before. The proxy for gender preferences GA_j is taken at the village/community level. The mean gender attitude for an individual i is calculated among adults born between 1951 and 1986 (about the same age as the parents of our target sample) for the village/community that child i resides in. One concern is that gender norms are effectively not randomly allocated across communities. Therefore, we also estimate the model with additional controls on communities' characteristics interacted with cohort dummies. We construct 15 communities' pre-policy characteristics from the village module of the CFPS-2010.[26] re Estimates controlling for communities' characteristics doesn't alter the results (see Appendix C, Table C2). As in the main analyses, we condition on province fixed effects (δ_p). Therefore, our empirical specification absorbs differences in child human capital across provinces and solely relies on within province community/village level variation in gender preferences. By conditioning on birth cohort fixed effects (γ_t), we aim to absorb all variation across age groups. We also include X_{ipt} to control for individual, family and provincial characteristics.

The main coefficient of interest is β_1, and, as we estimate the model separately by gender, the coefficient measures the within gender differential effect of the iodine fortification program across families with varying levels of son preferences. Mapping β_1 to the decomposition (4), the first term (B) on the right-hand side is the biological effect, which we assume to be homogeneous for a given gender and thus independent from parental gender preferences (α_s). As a consequence, β_1 will reflect the extent to which the behavioral response ($C \times D$) depends on α_s. We expect preferences to primarily have a role via D, the parental investment response to the salt-iodization program. The term C is the efficiency of the investment and depends on the health production technology. While C might be related to gender preferences α_s, it is expected that α_s plays a more direct role in D.[27] Note that the Gender Attitude variable is normalized with zero mean, and therefore the estimate of β_2 can be compared with the reform effect of section 5.

Table 5 reports the results by gender for cognitive outcomes. For females, high values of GA mean lower values of α_s in the theoretical model. The estimates of β_1 are reported in the first row of panel A (for males) and B (for females). The estimates of β_1 show large and significant effects (specifically for Math and Schooling) for females and (slightly) smaller and significant effects for males. Girls residing in villages/communities with strong preferences for boys benefit more from the universal salt iodization program than otherwise similar girls in communities/villages with less strong preferences for boys. The effects for boys is slightly smaller but with an opposite sign. Recall that high values of GA mean high values of α_s for boys in the theoretical model. Therefore, a negative estimate of β_1 suggests that boys from villages/communities with strong preferences for boys benefit benefit less from the universal salt iodization program. Estimates from both boys and girls suggest that the son preferences of parents drive at least part of the differential effects by gender of the effect of the salt iodization program on cognitive outcomes. Parental investments (D) in cognition in boys are partially crowded out by the universal salt-iodization program.

[26]The information in village module was collected from a knowledgeable individual who has access to statistical materials in the village, such as the director or the accountant of the community committee.

[27]The The efficiency of the investment may vary with gender preferences if, for instance, parents do not monitor the investment's effect. It is expected that an effect of α_s on C (if present) is a second-order effect.

Table 5: The Impact of Iodine Exposure by Gender Attitudes

	(1) Math Test ln(scores)	(2) Verbal Test ln(scores)	(3) Primary School	(4) Schooling ln(years)
Panel A: Males				
Post × Goiter × Gender Attitudes	-0.0617	-0.0748	-0.0613	0.00322
	[0.0335]*	[0.0535]	[0.0325]*	[0.0477]
Post × Gender Attitudes	-0.140	-0.143	-0.00764	-0.00815
	[0.137]	[0.123]	[0.0872]	[0.0886]
Post × Goiter	0.0533	0.0815	0.0311	0.0224
	[0.0498]	[0.0489]	[0.0247]	[0.0365]
Mean of Dep. Var.	2.517	3.083	0.497	2.107
Observations	3302	3302	3780	3644
Panel B: Females				
Post × Goiter × Gender Attitudes	0.176	0.0800	0.0333	0.0977
	[0.0558]***	[0.0409]*	[0.0249]	[0.0486]*
Post × Gender Attitudes	-0.0566	-0.0232	-0.0541	0.0620
	[0.175]	[0.129]	[0.144]	[0.261]
Post × Goiter	0.120	0.116	0.0342	0.0885
	[0.0391]***	[0.0536]**	[0.0103]***	[0.0220]***
Mean of Dep. Var.	2.543	3.171	0.529	2.122
Observations	3109	3109	3507	3383

Notes: Each coefficient is from a separate regression. All regressions control for fixed effects specific to birth province and birth year, birth order, family size, parents' education, region-specific linear trends, survey wave by age interactions and survey wave by cohort interactions. Mean-reversion controls include provincial average educational attainment in the 1990 Census interacted with the dummy for treated cohorts, hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2010 and CFPS-2014

One concern for our triple-difference results in Table 5 is that the gender selection could also drive these results. With boy biased gender preferences, parents could selectively abort female fetuses. In this case, the marginal girl born in a village with strong boy preferences may differ from the marginal girl in a gender-neutral village. We link the community/village level gender attitudes to sex ratio data to address this potential threat of endogenous sex selection. Ideally, we would like to construct sex ratio information at the community/village level. Unfortunately, this information is not available. Therefore, we rely on sex-ratio information at the city unit (the next level of spatial aggregation).[28] Like the event study design in section 5.4 we relate city level sex-ratios to gender attitudes by year, controlling for province fixed effects. The results of this event study are depicted in Figure 13. There

[28]There are 128 city units, each consisting of on average four communities/villages.

is no association between gender attitude and sex ratios, and, importantly, the association doesn't change after introducing the salt iodization policy. Combining the evidence of this analysis with the event study analyses in section 5.4, we can conclude that sex selection is not a likely driver of our findings in Table 5.

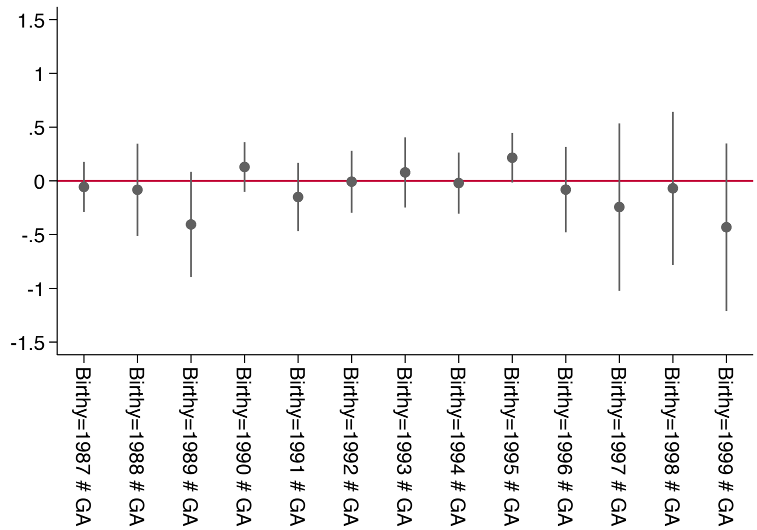

Figure 13. Sex Ratio and Gender Attitudes

Notes: Sex ratios are aggregated from microdata of the 2000 Census. Following Edlund et al. (2013), we calculate the sex ratio at city level for each birth year between 1987 and 1999. In the event study regression, we control for province fixed effects and birth year fixed effect. Standard errors are clustered by city.

The salt-iodization program did not affect parental investment in cognitive skills for boys. As argued before (section 5.4), parents could divert their investments in cognitive skills for boys to other dimensions of human capital, notably non-cognitive skills such as physical health, mental health, social skills, non-cognitive skills that enhance labor market outcomes, etc. Indeed, the results in Table 4 showed that boys do benefit along some non-cognitive skill dimensions. To see whether the program effects on non-cognitive factors vary with gender preferences, we also estimated the triple-difference model for the set of non-cognitive variables available for our cohorts, born between 1991-2000. This includes the subscores of the CES-D test and self-reported social skills. The estimates in the third row of Panel A (Boys) and B (Girls) in the Table C3 are the effects of the implementation of the salt iodization program (β_2) and are in line with the findings in Table 4. However, none of the effect estimates of β_1 in the first row of Panel A and B are significant for our set of non-cognitive skill measures, suggesting that the effect of the salt-iodization program does not vary by gender attitude. An explanation for this can be found in the primary impact of the salt-iodization program. In utero exposure to iodine deficiency has effects on fetal brain development primarily, and therefore, the reform's first order effects are on cognition. The reform also affects the allocation of resources over the different dimensions of human capital. This second-order effect is reflected in the estimates of the impact of the policy (β_2). Therefore, to what extend the impact of the policy on non-cognitive skills for boys (or girls) varies within gender by gender preferences (β_1) can be considered a third-order effect. Indeed, this effect might be small. It might be too demanding for the data and the measures of non-cognitive skills that we have at our disposal.

7 Conclusion

Currently, about 2 billion people suffer from iodine deficiency. The medical literature documents that iodine deficiency can lead to neurodevelopmental problems when fetuses are exposed in utero. This can lead to a reduction in children's cognitive skills and, consequently, adverse labor market outcomes later in life. This paper evaluates the effect of a nationally implemented salt iodization program on cognition of school-aged children in China. It differs from previous work as we explicitly focus on the role of gender preferences and how this may affect the effectiveness of large scale public programs. Gender preferences may also explain gender differences in the empirical literature on the long-run effects of adverse conditions early in life. Our difference-in-differences analyses find strong positive effects of the program for girls. A one standard deviation decrease (12%) in the pre-intervention children goiter rate is associated with math and vocabulary scores increasing by roughly 15%. We also see large increases in the educational attainment of females. Using the simple back of the envelope calculation, we infer that this translates to income increases of about 6,6%. Yet, we do not find any effects for boys.

These findings are robust against alternative specifications and falsification tests. These findings thus support the effectiveness of important, low costs, public health intervention. Indeed, compared to other interventions to raise education, the cost of salt iodization is extremely low. The costs associated with the intervention are about 0.05$ per person per year (WHO, 2005). This contrasts sharply with other interventions, such as class size reductions, costing over $5,000 (2010 dollars) per year per student (Chetty et al., 2011).

We proceed by further looking into these gender-specific findings and argue that gender preferences may play an important role. We do this in the context of a simple model along the lines of Yi et al. (2015); Almond, Currie, and Duque (2018) that explain how parental investments respond to (adverse) effects of shocks early in life. We show that gender preferences may lead to different investments in boys and girls. Following Yi et al. (2015); Almond, Currie, and Duque (2018) we use that total effect of an early life shock can be decomposed into a biological effect and a behavioral effect. The behavioral effect includes the parental investment decision, which mitigates or reinforces the biological effect. This behavioral ("resource allocation effect") depends on child preferences that may differ by gender. This simple model can explain the policy effect on girls and the absence of an effect on boys. Prior to the salt iodization policy, parents may have countered initial adverse shocks for boys and less so for girls. Therefore, when boy preferences are important, girls may benefit more from a nationally implemented programs. This also suggests that the salt-iodization program crowds out private parental investments in cognition. Another consequence of this may be that parents divert their investments into other skill dimensions, notably for boys. Indeed, we find program effects on non-cognitive skills for boys, but not for girls.

We test this hypothesis by explicitly accounting for gender preferences and proxy gender preferences with an index for gender attitudes. We estimate a triple-difference model per gender. Hence, the model identifies the effect of gender preferences from within gender variation in community/village gender attitudes. In line with the model and expectations, girls' program effects on cognition are stronger in communities/villages where preferences favoring boys are stronger. We do not find such effects for boys. Nor do we find such triple-difference effects for non-cognitive skills, both for boys and girls. Cultural or other contextual factors that feed attitudes favoring boys are not restricted to China but also hold for other countries in South-East Asia, the Middle East, and North Africa. The idea that gender preferences play a role in the western world can not be excluded either. Our findings, therefore, do not only speak to the external validity of the current study but also suggest that later life gender differences in labor market outcomes observed in many countries may be rooted early in life. The findings also indicate that generally large-scale programs may reduce gender inequalities and contribute to gender convergence.

Preferences favoring boys may lead to unequal investments in girls and justify policies that aim to reduce the consequences of such preferences. Also, improved economic circumstances may reduce such gender preferences, but as earlier work has shown (Almond and Edlund, 2008; Abrevaya, 2009) gender biases persist even with improved economic circumstances. This then calls for policies such as increasing mandatory schooling for both genders and other public programs that affect all, such as the public health campaign in this paper. Parental preferences are important, but this does not exclude the relevance of other factors such as financial constraints and the production technology of human capital formations. Indeed, for a significant part, these general investments can only improve outcomes as long as the investments translate into human capital; in terms of the model, the efficiency of investment (i.e., the term C in the decomposition (4)). With low values in the efficiency of investment, the impact of any investment (public or private) remains modest. The key to understanding the factors that drive the efficiency of the investment effect lies in understanding the human capital production technology. Indeed, this is one of the topics that require further development (see, e.g. Cunha, Heckman, and Schennach, 2010; Agostinelli and Wiswall, 2016; Attanasio, Meghir, and Nix, 2018). Unfortunately, with the data at hand, we can not competently tackle this problem and therefore leave this to future research.

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Appendix A Additional Information

In the analyses, we use a series of province-level pre-policy characteristics. We list all these variables here and describe how we construct each variable.

- Hospital per capita: Number of hospital per 1000 (Statistics yearbook 1991).
- Hospital beds per capita: Number of hospital beds per 1000 (Statistics yearbook 1991).
- School per capital: Number of primary schools per 1000 (Statistics yearbook 1991).
- Sex ratio: Sex ratio of the cohort 1991 calculated from Census 2000.
- Rural income: Average rural household income calculated from NFPS 1991.
- Poverty: Poverty rates in 1993 are obtained from Woo et al. (2004).

Table A1: Regional Classification of Provinces

Region	Provinces
North China	Beijing, Tianjin, Hebei, Shanxi and Inner Mongolia
Northeast China	Liaoning, Jilin and Heilongjiang
East China	Shanghai, Jiangsu, Zhejiang, Anhui, Fujian, Jiangxi and Shandong
South Central China	Henan, Hubei, Hunan, Guangdong, Guangxi and Hainan
Southwest China	Chongqing, Sichuan, Guizhou, Yunnan and Tibet
Northwest China	Shaanxi, Gansu, Qinghai, Ningxia and Xinjiang

Notes: Author's tabulations

Table A2: Summary Statistics

	High Goiter Provinces		Low Goiter Provinces	
	Females	Males	Females	Males
Outcomes				
Educational Attainment	2.02	1.94	2.04	1.95
	[0.79]	[0.79]	[0.83]	[0.79]
Illiterate	0.28	0.33	0.30	0.32
Primary School	0.46	0.42	0.40	0.44
Middle School	0.24	0.24	0.28	0.22
High School or above	0.028	0.014	0.029	0.024
Schooling	7.35	7.08	7.49	7.25
	[2.52]	[2.45]	[2.51]	[2.42]
Math Test Scores	12.2	11.9	12.5	12.2
	[5.23]	[5.25]	[5.13]	[4.79]
Verbal Test Scores	23.8	22.3	24.2	22.5
	[7.32]	[7.64]	[7.01]	[7.29]
Demographics				
Age	14.5	14.4	14.3	14.4
	[2.52]	[2.50]	[2.48]	[2.48]
Father's Educational Attainment	2.19	2.20	2.46	2.48
	[0.94]	[0.95]	[0.88]	[0.88]
Mother's Educational Attainment	1.68	1.69	2.16	2.13
	[0.85]	[0.83]	[0.92]	[0.88]
Birth Order	1.56	1.69	1.63	1.72
	[0.73]	[0.86]	[0.96]	[0.92]
Family Size	5.13	4.80	5.10	4.78
	[1.57]	[1.45]	[1.66]	[1.54]
Number of observations	956	1028	997	1007

Notes: Author's tabulations of CFPS-2010 and CFPS-2014. Sample consists individuals born in rural area between July 1995 and June 2000. We label a province as high/low goiter if its goiter rate is above/below 17% (median).

Appendix B Additional Results: Iodine and Long-Run Outcomes

Table B1: Robustness Checks (Male)

	(1) Math Test ln(scores)	(2) Verbal Test ln(scores)	(3) Primary School	(4) Schooling ln(years)
Panel A: raw test scores				
Post × Goiter	0.108	0.405	0.0202	0.0878
	[0.333]	[0.473]	[0.0233]	[0.169]
Mean of Dep. Var.	13.59	23.72	0.499	8.569
Observations	3310	3310	3803	3654
Panel B: additional controls				
Post × Goiter	-0.0588	-0.0580	-0.00554	-0.0143
	[0.0507]	[0.0538]	[0.0297]	[0.0686]
Mean of Dep. Var.	2.517	3.083	0.499	2.108
Observations	3310	3310	3803	3654
Panel C: drop partial exposed group				
Post × Goiter	0.0296	0.0575	0.0260	0.0216
	[0.0517]	[0.0524]	[0.0230]	[0.0403]
Mean of Dep. Var.	2.514	3.081	0.497	2.106
Observations	3235	3235	3706	3560
Panel D: only using baseline wave 2010				
Post × Goiter	0.0551	0.0809	-0.00890	0.0111
	[0.0567]	[0.0666]	[0.0290]	[0.0369]
Mean of Dep. Var.	2.486	3.046	0.344	1.978
Observations	1992	1992	2028	2028
Panel E: small sample window				
Post × Goiter	0.0633	0.0668	0.0223	0.0344
	[0.0584]	[0.0660]	[0.0213]	[0.0465]
Mean of Dep. Var.	2.549	3.108	0.516	2.129
Observations	2643	2643	3047	2920

Notes: Each coefficient is from a separate regression. All regressions except Panel B use the same controls as the baseline model in Table 2. In Panel B, we control for birth-region and birth-year specific interaction instead, and we additionally control for schools per capital in 1991, poverty rates in 1993, and average household income in 1991 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2010 and CFPS-2014

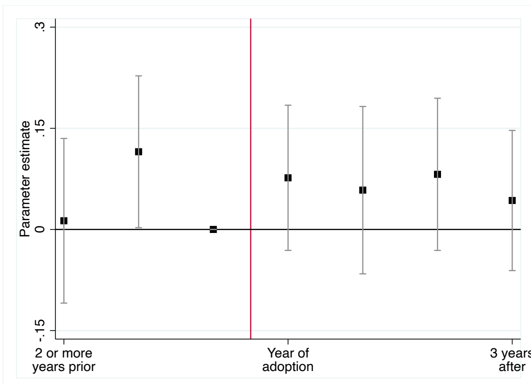

Figure B1. Math Test Scores

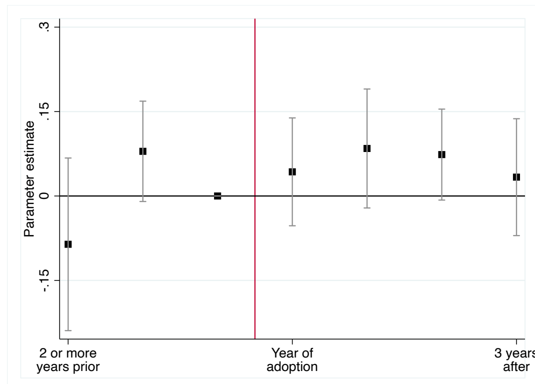

Figure B2. Verbal Test Scores

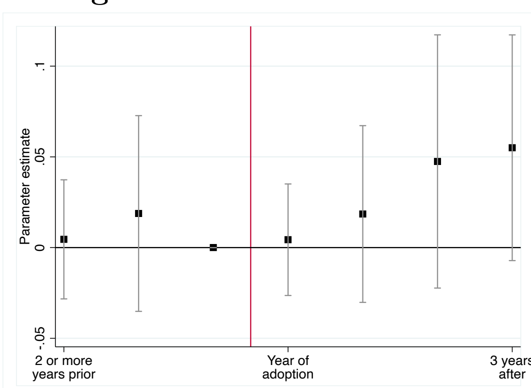

Figure B3. Primary School

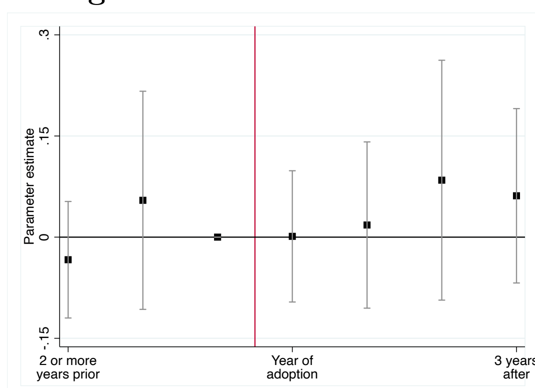

Figure B4. Schooling

Notes: The sample includes all male respondents from two waves of the survey (CFPS-2010 and CFPS-2014). Each point reflects the coefficient estimated on an interaction term between the birth year (compared to 1995) and the pre-intervention (base) level of the goiter rate in the birth-province. Capped spikes represent 95 percent confidence intervals. All models condition upon birth province and birth year fixed effects and the full set of controls used in our main analysis.

Appendix C Additional Results: More on Gender

C.1 Iodine Exposure and Non-Cognitive Skills

CES-D includes 20 questions, which were aggregate to four categories: somatic complaints (Q1, Q2, Q5, Q7, Q11, Q13, Q20); depressed affect (Q3, Q6, Q9, Q10, Q14, Q17, Q18); positive affect (Q4, Q8, Q12, Q16)[29]; interpersonal problems (Q15, Q19).

Questions: Below is a list of the ways you might have felt or behaved. Please tell me how often you have felt this way during the past week with a score: 1 for rarely or none of the time (less than 1 day); 2 for some or a little of the time (1-2 days); 3 for occasionally or a moderate amount of time (3-4 days); 4 for most or all of the time (5-7 days).

1. I was bothered by things that usually don't bother me.
2. I did not feel like eating; my appetite was poor.
3. I felt that I could not shake off the blues even with help from my family or friends.
4. I felt I was just as good as other people.
5. I had trouble keeping my mind on what I was doing.
6. I felt depressed.
7. I felt that everything I did was an effort
8. I felt hopeful about the future.
9. I thought my life had been a failure.
10. I felt fearful.
11. My sleep was restless.
12. I was happy.
13. I talked less than usual.
14. I felt lonely.
15. People were unfriendly.
16. I enjoyed life.
17. I had crying spells.
18. I felt sad.
19. I felt that people dislike me.
20. I could not get "going".

Table C1: Summary Statistics of Gender Attitudes

	Males	Females
Son should live together with his parents.	3.49	3.47
	[1.34]	[1.38]
Every family should at least have a son.	3.46	3.36
	[1.48]	[1.55]
The husband takes care of the business, and the wife takes care of the family.	4.08	4.07
	[1.10]	[1.15]
Woman's marriage is more important than her career.	3.50	3.75
	[1.33]	[1.27]
Every woman should have a child.	4.08	4.33
	[1.17]	[1.02]
Disagree: The husband should do half of the housework.	1.95	1.84
	[1.10]	[1.05]
Gender index	3.43	3.47
	[0.70]	[0.71]
Normalized gender index	-0.031	0.030
	[0.99]	[1.01]
Number of observations	10070	10455

Notes: Sample includes all individuals born between 1951 and 1986 in CFPS-2014. Surveyed respondents were asked if they agree with these six statements. The respondents report how much they agreed with a certain statement on a scale of 1-5, with 1 being *Strongly agree* and 5 being *Strongly disagree*. Gender index is the average of the 6 indicators. Normalized gender index is calculated by subtracting the mean and dividing by the standard deviation.

C.2 Additional Results on Gender Attitudes

To show that parental attitudes reflect parental preferences is related to parental investment in children, we exploit the fact that CFPS does include some information about parental investments behavior for a different sample of children younger than 6 years of age. For this sample we construct a "parental investments index" by calculating an unweighted average of the z-score of 5 variables: breastfeeding practices; whether the child went to kindergarten; how often the parent read to their child; how often the parent buys books for their child; and how often they travel with their child.[30] Breastfeeding is a dichotomous variable whether the duration of breastfeeding was longer than three months. Kindergarten is an dichotomous variable indicating whether the child has attended kindergarten. The last three variables are categorical variables for the frequency of the event ranging from 0 ("never") to 5 ("every

[29]We reverse the score of each question in the positive affect according to the literature (i.e., 1 means most or all of the time (5-7 days) and 4 means rarely or none of the time (less than 1 day); 2 for occasionally or a moderate amount of time (3-4 days); 3 for some or a little of the time (1-2 days)).

[30]Iodine could also be obtained from food such as vegetables, eggs, fish and meat. The composition of the diet as well as the quantity of food may therefore influence the iodine intake of pregnant women and infants. The eating of healthy and nutritionally rich food could be viewed as an antenatal investment and may differ with knowledge of the gender of the unborn child. Unfortunately we do not have information on food consumption of pregnant women and infants.

Table C2: The Impact of Iodine Exposure by Gender Attitudes

	(1) Math Test ln(scores)	(2) Verbal Test ln(scores)	(3) Primary School	(4) Schooling ln(years)
Panel A: Males				
Post × Goiter × Gender Attitudes	-0.0757	-0.102	-0.0518	-0.0284
	[0.0430]*	[0.0607]	[0.0423]	[0.0447]
Post × Gender Attitudes	-0.0931	-0.0180	0.105	0.105
	[0.153]	[0.110]	[0.0928]	[0.138]
Post × Goiter	0.0310	0.0689	0.0235	0.0112
	[0.0475]	[0.0445]	[0.0227]	[0.0360]
Mean of Dep. Var.	2.517	3.082	0.497	2.107
Observations	3257	3257	3731	3598
Panel B: Females				
Post × Goiter × Gender Attitudes	0.136	0.0209	0.0517	0.0750
	[0.0671]*	[0.0599]	[0.0301]*	[0.0433]*
Post × Gender Attitudes	0.0124	0.0644	-0.0741	0.171
	[0.217]	[0.157]	[0.137]	[0.284]
Post × Goiter	0.110	0.125	0.0307	0.0857
	[0.0339]***	[0.0499]**	[0.0124]**	[0.0208]***
Mean of Dep. Var.	2.541	3.171	0.528	2.122
Observations	3062	3062	3457	3334

Notes: Each coefficient is from a separate regression. All regressions control for fixed effects specific to birth province and birth year, birth order, family size, parents' education, region-specific linear trends, survey wave by age interactions and survey wave by cohort interactions. Mean-reversion controls include provincial average educational attainment in the 1990 Census interacted with the dummy for treated cohorts, hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990 all interacted with cohort dummies. Additionally, communities controls include cohort dummies interacted with 15 communities' pre-policy (1991) characteristics: access to electricity, radio, satellite TV, post services, phone, mobile phone, highway, rail, tap water, gas, local factory, hospital, local election, urbanization, and land property rights. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2010 and CFPS-2014

day"). We take higher values of the score to be associated with more investments in the child. It is not obvious that this also holds for the kindergarten indicator. Recent papers on the effect of day care (children aged 0-2) find mixed results. Positive effects are found for Norway (Drange and Havnes (2019) shown significant gains in language and mathematics at age 6–7 of childcare enrollment using childcare assignment lotteries.) and Germany (Felfe and Lalive (2018) shown early child care has a strong positive effects on children's motor and socio-emotional skills.) while the recent paper by Ichino, Fort, and Zanella (2019) find negative effects, in particular for girls.[31] Note that kindergarten usually concerns older infants (age 4

[31] They examine the effects of extended day care of children aged 0-2 on cognitive and non-cognitive skills. The idea is that daycare implies fewer one-to-one interactions with adults as inputs in the technology of skill

Table C3: Iodine Exposure and Non-Cognitive Skills

	(1) Somatic Complaints	(2) Depressed Affect	(3) Positive Affect	(4) Interpersonal Problems	(5) Social Skills
Panel A: Males					
Post × Goiter × Gender Attitudes	0.143	-0.288	-0.680	-0.139	0.737
	[0.243]	[0.331]	[0.292]**	[0.119]	[0.649]
Post × Gender Attitudes	-0.740	-0.655	0.287	-0.0193	0.320
	[1.267]	[0.971]	[1.534]	[0.452]	[2.525]
Post × Goiter	-0.487	-0.438	0.121	-0.0852	0.710
	[0.197]**	[0.166]**	[0.168]	[0.0496]*	[0.346]*
Mean of Dep. Var.	3.340	2.587	4.808	0.576	22.29
Observations	1554	1556	1556	1555	1064
Panel B: Females					
Post × Goiter × Gender Attitudes	-0.559	-0.289	-0.129	-0.245	0.495
	[0.418]	[0.365]	[0.325]	[0.182]	[0.835]
Post × Gender Attitudes	-0.415	0.519	0.107	0.544	-0.00960
	[1.305]	[1.105]	[0.902]	[0.407]	[2.011]
Post × Goiter	-0.0797	-0.114	0.0518	-0.00230	-0.180
	[0.217]	[0.299]	[0.172]	[0.0726]	[0.350]
Mean of Dep. Var.	3.515	3.259	4.872	0.653	22.43
Observations	1436	1436	1435	1436	1017

Notes: Each coefficient is from a separate regression. All regressions control for fixed effects specific to birth province and birth year, birth order, family size, parents' education, region-specific linear trends and age. Mean-reversion controls include hospitals per capita in 1991, hospital beds per capita in 1991, and the sex ratio in Census1990 all interacted with cohort dummies. Standard errors clustered by province appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2012 and CFPS-2014.

and older). Also, for our sample of rural families interaction with other infants may be more beneficial and can be viewed as an investment.

For this sample of parents who have children younger than 6, we run a regression that relates the parental investment index to our proxy for gender preferences:

$$I_{ij} = \beta_0 + \beta_1 Male_i \times GA_j + \beta_2 Male_i + \beta_3 GA_j + X_i \rho + \epsilon_i, \tag{6}$$

where I_{ij} is parental investments index for child i from village/community j. $Male_i$ is dummy variable indicating the gender of the child, and GA_j is the gender attitude at the

formation and therefore be more harmful for infants. Girls at age 0–2 are relatively more capable of making good use of stimuli that improve their skills. Extended exposure to daycare may therefore be particularly harmful for girls.

village/community j.[32] Similar to Dossi et al. (2019), we also control for a series of variables X_i such as parental education, birth order, family size and province fixed effects. Prime interest is in the parameter β_1, the association between gender preferences and parental investment in young children.

The results of the regression are reported in Table 5. In first three columns, we gradually add family background controls to the regression. In the fourth column, we also control for province fixed effects. In this specification the estimates solely rely on within province variation in gender attitudes. The estimates of β_1 are reported in the first row of Table 5 and show that gender attitudes are strongly associated with parental investments in children and that investments in girls are lower in areas where preferences favoring boys are stronger. Adding controls does not affect the estimate of β_1. The small change in β_1 when we add province fixed effects (column 4) also underlines that there is substantial within province variation in the gender attitude variable. This is important for the section 6.3 where we exploit this variation to identify the role of preferences in the effect of the iodization policy on cognitive and non-cognitive outcomes.

Table C4: Gender Attitudes and Parental Investment

	(1)	(2)	(3)	(4)
	Parental Investment Index			
Gender Attitudes × Male	0.163	0.189	0.186	0.165
	[0.0834]*	[0.0816]**	[0.0824]**	[0.0846]*
Male	-0.0665	-0.0495	-0.0489	-0.0381
	[0.0329]**	[0.0325]	[0.0321]	[0.0330]
Gender Attitudes	-0.315	-0.248	-0.221	-0.170
	[0.0771]***	[0.0711]***	[0.0728]***	[0.0791]**
Parental Education	No	Yes	Yes	Yes
Birth Order & Family Size	No	No	Yes	Yes
Province Fixed Effects	No	No	No	Yes
Observations	1,642	1,642	1,642	1,642

Notes: Each column is from a separate regression. Standard errors clustered by village/community appear in square brackets. *, **, *** indicates significance at the 10%, 5% and 1% level respectively.
Data: CFPS-2010, CFPS-2012 and CFPS-2014

[32] With this regression we do not aim to make causal statements. Nevertheless, in the calculation of the village/community level average we left out the individual response to minimize confounding.