Self-Reinforcing Gender Equality from Labor Market to Child Health

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Abstract: Achieving gender equality is a crucial development goal worldwide and can also serve as a key driver of broader social progress. But can promoting gender equality trigger a virtuous circle, or become self-reinforcing? We propose and test a model of the self-reinforcing effect of gender equality, in which improvements in the labor market for the parent generation promote human capital for the next generation of children, especially for girls. Using data from China, we find that a one-standard-deviation increase in the female-to-male income ratio — driven by exogenous changes in the labor market—leads to a direct increase of a 0.032-point in children's height (in standard deviation terms), which is equivalent to an average height gain of 0.48 cm among seven-year-old children. The positive effect is stronger for girls, who are 0.029 point more affected than boys, indicating that gender equality can reproduce and strengthen itself over time, with dynamic and long-term impacts. We further explore how this self-reinforcing mechanism of gender equality varies across different social environments. The effect is found to be significant only in contexts with more modern living conditions, higher parental labor market participation, and weaker son-preference norms.

Keywords: Gender equality; Income gap; Child health; Self-Reinforcing effect

JEL classification: J16; I14; J31; O15

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Gender inequality is the overwhelming injustice of our age and the biggest human rights challenge we face. But gender equality offers solutions to some of the most intractable problems of our age.

——António Guterres，UN Secretary-General，02 March 2020

1. Introduction

Equality serves as an important goal and driving force in the development of human society (Rousseau, 1754) , with gender being one of the most important dimensions. It is widely acknowledged that a more gender-equitable environment is more inclusive and harmonious, which is conducive to socioeconomic development (UN, 2020). In other words, gender equality can promote economic development. However, most studies have shown that conversely, economic development itself does not necessarily guarantee steady progress in gender equality (Eastin and Prakash, 2013; Duflo, 2012). As Duflo (2012) put it, "Does this imply that pushing just one of these two levers—'women empowerment' and 'economic development'—would set a virtuous circle in motion?" Her answer is that the existing literature suggests the interrelationships between them may be too weak to be self-sustaining.

Existing literature has primarily treated gender equality as a slowly changing outcome variable, determined by long-term historical, economic, and cultural factors (Jayachandran, 2015). Deep historical forces and persistent social norms are often highlighted, for example, long-ago agricultural practices continue to influence modern gender roles (Alesina et al., 2013), and regions with matrilineal traditions or other historical institutions show enduring differences in gender outcomes (Bazzi et al., 2023). Economic structure and growth patterns also matter. Qian (2008) demonstrates that sectoral economic booms can affect the female–male balance, and Guiso and Zaccaria (2023) link changing industry composition to evolving gender gaps. Cultural preferences and technological changes further condition gender inequality (Afridi et al., 2023). These studies, however, largely view gender equality as a static equilibrium determined by external conditions. They leave open the possibility of a dynamic feedback mechanism through which gender equality might reinforce itself across generations.

Some research has focused on the impacts of gender equality. It has established a positive correlation between women's economic empowerment and improvements in child well-being, particularly from the perspective of labor market dynamics and intra-household decision-making. Recent syntheses emphasize that the most compelling evidence for these links typically comes from settings in which women's earning capacity or control rights change in a plausibly exogenous manner, because such changes can shift intrahousehold power and reallocate resources toward children. This evidence is generally more informative than simple correlations between empowerment proxies and child outcomes (Jayachandran and Voena, 2026). Duflo (2003) shows that pensions paid to South African grandmothers, but not grandfathers, increase allocations benefiting granddaughters, underscoring that who receives income matters for child human capital. Newer evidence such as Calvi (2020) exploits changes in inheritance law in India to demonstrate that institutional shifts that bolster women's claims can raise their resource shares and reduce poverty. These indicate that who earns or controls income within a household can profoundly influence child outcomes (Lundberg et al., 1996). At the macro level, greater female labor force participation and earning power have been linked to improved child health and education, operating through higher household incomes and more equitable resource allocation. In settings where women gain economic agency, families tend to spend more on nutrition, schooling, and healthcare for their children (Blau & Kahn, 2017).

The labor market is not only the primary channel for people to earn income but also a key area where progress in gender equality has been made in developing countries. On one hand, persistent gender discrimination in the labor market constrains women's economic agency, weakening their ability to negotiate favorable outcomes within the household (Bertrand et al., 2010). By contrast, gender-biased shifts in labor market conditions that favor women can alter intra-household decisions in ways that benefit children, especially girls. Qian (2008) shows that the increase in women's farm income (driven by a boom in tea prices) led to higher survival rates for infant girls and greater educational investment in daughters in rural China, whereas increases in men's income (e.g., from orchard crops) had the opposite effect. Jensen (2012) finds that introducing new job opportunities for young women through the business process outsourcing industry induced families to invest more in girls' education and improved their health and nutrition. Similarly, evidence from

Bangladesh shows that the rapid expansion of garment factory jobs has led to a 50% increase in school enrolment among girls in villages near new factories, along with delayed marriage and childbearing (Heath & Mobarak, 2015). These studies demonstrate that when women's expected earnings increase, parents respond by reallocating resources toward daughters. There are also empirical studies based on China ' s newly and rapidly changing labor market that have linked such gender equality to parental investments in children. Wang and Cheng (2021) examined the impact of the gender wage gap on household educational expenditures for children in urban China. They found that reductions in the gender wage gap are associated with increased household investment in children's education, with a slightly more pronounced effect observed for male children. Their findings highlight that household income and female bargaining power partially mediate the relationship between gender wage disparities and parental investment in children's schooling. Similarly, Zhao and Qu (2024) studied the effects of regional economic zone policies on intra-household bargaining power in rural China, demonstrating that increased female bargaining power within the household positively correlates with greater human capital accumulation among children.

These studies suggest that improvements in women's relative economic position can translate into gains for the next generation. However, prior studies have mostly examined specific policy shocks or local changes, and they often emphasize mechanisms like maternal altruism or increased bargaining power from current income. Less explored is the idea of a forward-looking, self-reinforcing dynamic in which expectations of women's future opportunities drive sustained changes in household behavior. Moreover, while several literatures identify overall gains for children, the conditions under which sons or daughters benefit differentially remain a question shaped by local labor-market structure and social norms.

Building on the aforementioned findings, this study contributes new evidence that gains in gender equality in one generation can propagate intergenerationally, reinforcing equality in the next generation. In contrast to explanations based on static bargaining models, our framework highlights forward-looking household decision-making driven by "expected returns". Parents, as rational economic agents, invest in their children's human capital with an eye to future payoffs. When the labor

market signals that daughters will face better opportunities and higher earnings, the perceived return on investing in girls' health and education increases. In response, parents optimally allocate more resources to their daughters. This expected-return channel alters the household's optimal investment rule by shifting intrahousehold Pareto weights and increasing the shadow value of resources allocated to girls (Manser and Brown, 1980; Lundberg and Pollak, 1993, 1996; Jayachandran and Voena, 2026). By framing parental decisions as forward-looking investments, this mechanism offers a strategic and investment-oriented explanation for reductions in bias against daughters, moving beyond interpretations that attribute improved outcomes solely to greater maternal altruism or increased female decision-making authority.

Over time, this process can become a virtuous cycle – essentially an economic engine that helps perpetuate and even accelerate gender equality across generations. Specifically, we explore how macro-level improvements in gender equality in the labor market translate into more equitable investment in human capital at the micro-level of households, thereby fostering an environment conducive to sustained gender equality. Theoretically, this aligns with Qian (2008)'s framework, emphasizing the importance of enhancing the social value of women and improving girls' human capital as key strategies for achieving gender balance and mitigating son preference. Moving beyond extreme survival conditions, this study focuses on children's—particularly girls'—health outcomes, exploring how intra-household income dynamics influence resource allocation.

Children's height is widely recognized in economics and development studies as a key indicator of early-life health, nutrition, and human capital. Some literature has established a causal link between early-life health, proxied by height, and long-term economic success, operating primarily through the channel of cognitive development (Case and Paxson, 2008; von Hinke et al., 2013). Unlike self-reported indicators such as food consumption, height is an objectively measurable trait that reduces the risk of misreporting or bias in economic studies. This makes height a reliable proxy for child well-being, particularly in low-income settings. For example, Jayachandran and Pande (2017) document that India's extreme child stunting reveals deep gender bias. Widely used indicators of educational attainment may be influenced by cultural biases and

variations in educational systems, making them less reliable for capturing short-term differences in human capital investment. Moreover, compulsory education laws may often obscure hidden gender disparities in childhood human capital investments. In contrast, height may serve as a more suitable measure for assessing immediate human capital conditions. Its universality enables consistent comparisons across different populations and settings, making height a valuable metric for studying disparities in early-life development and well-being.

The biologically expected pattern of growth trajectories is characterized by girls experiencing an early growth spurt (particularly between ages 10–13), after which boys gradually catch up and eventually surpass girls during adolescence (Abbassi, 1998). Such growth patterns have been observed in the Philippines, an Asian country without strong son preference. By contrast, in China, where son preference is prevalent, boys consistently maintain a greater height advantage over girls throughout the entire growth period, and this advantage is significantly larger compared to the male–female height difference observed in the Philippines (Song & Burgard, 2008). The data used in this paper come from the China Family Panel Survey (CFPS), and the corresponding evidence can be found in the data. After ranking the provinces from worst to best female-to-male income ratios, Figure 1 presents three provinces at the 10th, 50th, and 90th percentiles to observe the growth trajectories. In provinces at the 10th and 50th percentiles, boys remain consistently taller than girls throughout childhood. Only in the provinces at the 90th percentile, growth trajectories align with the biologically expected pattern, with girls initially surpassing boys in early adolescence before eventually being overtaken.

Insert Figure 1

The primary contribution of this study is to establish a causal relationship between gender equality in the labor market and its intergenerational transmission to children's health, thereby revealing the self-reinforcing nature and mechanisms of gender equality. To the best of our knowledge, this is the first paper to emphasize and confirm that gender equality can be self-reinforcing, as well as one of the rare pieces of optimistic empirical evidence. Health indicators, although often overlooked in

existing research, reflect outcomes beyond purely monetary investments and provide substantial, visible returns in the future, making them particularly salient for both parents and children. This study aims to present a more complete chain of the mechanism through which gender equality reproduces itself, emphasizing expected returns (or potential market value) as key drivers shaping household investments in human capital. The second contribution is to test under what social circumstances the positive reproduction effect works. The magnitude of the effect is expected to differ based on household characteristics and social contexts. Specifically, the study investigates whether the health effects differ between boys and girls, and further analyzes variations between urban and rural households, as well as among families with varying levels of traditional clan culture. Third, we use the gender difference in employment growth rate as the "Bartik instrument" to measure the wage gap at the county level, helps further address concerns about endogeneity caused by potential reverse causality in the regression. Moreover, we employ both anthropometric and clinical indicators to assess children's health outcomes, making the measurement more thorough. Through these aspects, this study aims to provide insights into the relationship between gender equality in China's labor market and child health outcomes.

The remainder of this chapter is structured as follows. Section 2 presents the background and a review of the theoretical framework. Section 3 describes the data set used in the regression and outlines the models. Sections 4 and 5 report the main results and a series of robustness tests. Finally, we conclude in section 6.

2. Background and Theoretical Framework

Background

The gender wage gap represents a substantial equality issue that affects labor market efficiency, as well as the allocation of household resources. China’s rapid development in recent decades has been accompanied by significant fluctuations and regional variation in gender equality, especially in the labor market. Figure 2a, based on CFPS data, illustrates a persistent income gap. A female-to-male income ratio equal to 1 means equal wages for men and women. The nationwide income ratio rose from 0.66 in 2010 to 0.75 in 2016, before falling back to 0.66 in 2018. Compared with

the world average of 0.77, this represents a large gender income gap. Figure 2b shows the relationship between provincial female-to-male income ratios and GDP per capita. There are wide variations in income ratios across regions, for example, ranging from 0.57 in Anhui to 0.84 in Tianjin. Over the period from 2010 to 2018, the top five economically well-developed regions, which have high levels of income equality, stand out. However, in most of the remaining regions, the income equality level is not strongly related to the level of economic development (Alesina et al., 2013; Duflo, 2012). This provides a good empirical background for discussing the transmission of equality in the labor market and within households.

Insert Figure 2

Children's health is our main concern. It is not only bound by income but also influenced by parents' decisions on family resource allocation. Household decisions in China have outsized importance for child outcomes because public social safety nets for children's health and welfare remain limited in many areas. Families are the primary providers of children's health care and nutrition, meaning shifts in household resource allocation can have immediate effects on child well-being. We propose a research question of whether there is a causal relationship between gender income equality and children's health, specifically height. Figure 3 shows the distribution of the average child height-for-age-gender z-score (HAZ)[3] by province in China. Provinces are ranked by gender income ratio from lowest to highest, with a trend of children's HAZ moving to the right, i.e., increasing, as the income gap narrows. We hypothesize that gender income equality may partially explain variation in children's height. The causal components of this relationship will be tested through empirical analysis and to elucidate the mechanism behind it.

Insert Figure 3

Theoretical Framework

[3] Further details are provided in later sections.

We propose a conceptual framework in which parents allocate household resources for children's health and education based on the expected future returns to those investments. In traditional models of intra-household allocation (Becker, 1981), if daughters have lower expected earnings, parents may rationally invest less in girls' human capital. However, as women's expected income rises, for instance, through better education or job opportunities, the perceived returns to investing in daughters increase. Our model posits that forward-looking parents will respond by raising both the total investment in children and the share of that investment going to daughters since the gender gap in returns has shrunk. This is not merely about current income effects but about a reallocation driven by changing expectations. The mechanism operates even if parents have no inherent gender bias or altruistic favoritism; it emerges from optimizing behavior under changing constraints. This focus captures the more continuous decision-making dynamics of nutritional status, which ultimately determines whether a child can thrive and later capitalize on educational opportunities.

In our framework, even if the initial allocation of power or altruism is unchanged, fathers and mothers alike have an incentive to invest more in daughters when those daughters' future prospects improve. In essence, the household's objective function places weight on children's future welfare (Lundberg & Pollak, 1993; Manser & Brown, 1980), and higher expected returns to investing in girls raise the marginal utility of allocating resources to them today. In fact, we acknowledge a complementary bargaining-power channel that unfolds over time, but this mechanism can operate before any shift in bargaining power occurs. It provides an economic incentive-based explanation for improving gender equality. Parents are not necessarily becoming more enlightened or egalitarian in preferences, but they respond to market signals that investing in daughters "pays off." It also implies that gains for daughters need not come at the expense of sons in absolute terms. If the total future returns increase by investing in all children, human capital investments can rise, though relatively more goes to daughters until an optimal balance is reached. Such a mechanism may help explain changes in behavior even in societies where traditional attitudes are slow to change.

The above mechanism produces a dynamic, intergenerational feedback loop – a

virtuous cycle in which gender equality in one generation seeds further equality in the next. The process can be thought of as a virtuous upward spiral, akin to what Duflo (2012) describes as the synergy between women's empowerment and economic development. Our emphasis is on the micro-level household investment decisions as the transmission mechanism for that synergy. It is important to note that such a positive feedback loop may not initiate or sustain itself under all conditions; severe discriminatory norms or institutional barriers can stifle the transmission. But when the loop does take hold – as we hypothesize is possible in transitional contexts like contemporary China – it offers a powerful explanation for how gender equality can be not just the result of development, but also a driver of further development, creating a self-perpetuating trajectory toward greater equality and better overall social outcomes. Greater gender equality (in terms of opportunities and incomes) today lays the groundwork for more rapid gains in women's status tomorrow, by altering familial decisions in favor of girls. Moreover, there are direct intergenerational externalities to women's human capital that fortify this loop (Schultz, 2002). Over time, this process can help break down traditional biases, for instance, if educated mothers demonstrate the productivity of daughters, societal preferences may shift away from son bias.

3. Data and Methods

3.1 Data

This paper adopts household-year level panel data from five waves of the China Family Panel Survey (CFPS) in 2010–2018. The selected household sample consists of married couples with children aged between 0 and 16. There are a few advantages to using the CFPS data. The survey is household-based and contains information on the health of the child generation and the labor market outcomes of the parental generation. It is a large-scale household tracking survey in China, featuring a broad group of respondents. Information on a series of family and personal characteristics in the CFPS allows for better control of variations in economic status, lifestyles, and other relevant factors.

Children health

The dependent variables are derived from the child questionnaire, using the height-for-age-gender z-score (HAZ) and the number of sicknesses in the past month. HAZ is a widely used indicator for assessing growth retardation (chronic malnutrition) in children, allowing for comparisons across gender and age groups. It is calculated as: $HAZ_{ag} = (Height_{ag} - Height_{ag;mean})/SE(Height_{ag})$, where '$a$' and '$g$' correspond to the age and gender group of the child. It captures the differences in growth patterns and deviations within the Chinese population. Our observed children's HAZ follows a normal distribution.[4] Moreover, following (Jakaria et al., 2022; Liu et al., 2024), the binary variable "Stunting" is identified by HAZ<-2, to be used in the robustness tests.

Apart from height information, the "number of sicknesses in the past month" is selected to measure short-term sickness status in the alternative test. This information is reported by the parents and reflects how many times the child has been sick in the past month. Using monthly sickness counts helps avoid measurement errors that could arise from inaccurate recall when using yearly sickness counts. Counting the number of sicknesses rather than the number of hospital visits mitigates bias stemming from different caregiving approaches to the child's health problem. The average frequency of sickness in a month in the sample is 0.520, and it is similar for both genders (0.533 for boys vs. 0.505 for girls).

Gender income gap

Labor market equality is assessed through the gender income ratio at the county level. Income data are sourced from the CFPS adult questionnaire, from respondents who reported being "at work in the last 12 months." The "total income from work" is aggregated at the county level by gender. The gender income gap is calculated as the ratio of female-to-male income in each county where the household resides. Three methods are used to measure income. First is the "general mean income". The county-level gender-specific labor income is calculated by actual mean earnings. Second is the "demographic-based income". Each sample is assigned to a demographic group based on gender, education level, and county, generating labor

[4] To eliminate the outliers and mistranscription, samples with HAZ less than -5 and greater than 5 are removed.

income at the demographic group level. Third is the "Bartik instrument". We construct a Bartik instrument using employment growth to enhance exogeneity and improve the accuracy of the estimation.

Differences in the distribution of occupations are an important factor in the existence of the gender wage gap (Blau et al., 2017), and there are two potential risks in using the general gender income gap. First, at the household level, income disparities may be influenced by the intra-household division of work and occupational choices. This could lead to concerns about reverse causality between child health care and occupational choices. For example, an increase in childcare responsibilities within the family might prompt parents, especially mothers, to choose more flexible work or shorter working hours, which may not accurately reflect their actual labor productivity or potential earnings. Second, at the regional level, the focus of industrial development can vary across different regional administrations. Factors such as policy support from central or local governments, or external industry shocks, can lead to differences in earnings across regions. A region's industrial composition can also influence workers' occupational choices, which in turn affects their income. For instance, regions with a strong focus on mining and construction industries might offer more employment opportunities for male labor, while regions with a significant presence in garment, electronic manufacturing, and service industries might see increased opportunities for female labor.

Following the methodologies of Goldsmith-Pinkham et al. (2020) and Akesaka & Kikuchi (2024), we construct gender-specific employment growth differentials as a Bartik instrumental variable. We begin by adopting the approach proposed by Bartik (1991), which is based on the idea of using national employment growth rates across industries, weighted by local industry employment structures, to derive gender-specific employment growth rates. The "Bartik instrument" is calculated as follows:

$$\overline{D_{pt}^{g}} = \sum_{j} r_{jp,2010}^{g} * E_{jt}$$

where D_{pt}^{g} denotes the Bartik employment growth rate in province p, year t for gender group g. The "share" term $r_{ejc,2010}^{g}$ represents the proportion of employment in industry j for gender g in province p as of the base year 2010.

These employment shares are obtained from the 2010 Population Census and cover 19 industrial classifications that remain consistent throughout the study period.[5] The shift term E_{jt} is the national employment growth rate in industry j in year t, derived from the China Statistical Yearbook. We then define the gender employment growth differential as: $\overline{Z_{pt}} = \overline{D_{pt}^{male}} - \overline{D_{pt}^{female}}$, which reflects the percentage point difference in employment demand growth between males and females.

Theoretically, the shift-share approach is exogenous and correlated with the gender income ratio. Within the standard Beckerian framework, relative wages adjust in response to relative labor demand shocks. The national industry-level growth rates (the "shifts") are considered exogenous shocks, as they are not expected to be influenced by the economic conditions of any single region, nor are they associated with unobserved regional characteristics. The "shares" are based on employment structures from the 2010 census, predating the study period, and thus are unaffected by subsequent child health investments or other potential confounders. These initial shares can therefore be regarded as exogenous. In addition, using the employment growth difference rather than a ratio of shocks keeps the instrument linear in levels, which simplifies the first-stage specification and avoids heteroskedasticity and the potential explosion of variance that arise when male employment growth is small or negative.

Figure A1 illustrates the base-period employment shares of females and males across provinces by industry category. If a particular industry contributes equally to male and female employment, the corresponding provincial data points would lie along the 45-degree line. We find a gender-disproportional contribution to local employment across industries. Additionally, the relative employment shares of each industry vary significantly across provinces. These observations suggest independent fluctuations in employment opportunities for males and females, thereby supporting

[5] The 19 classifications include: (1) Agriculture, Forestry, Animal Husbandry, and Fishing; (2) Mining; (3) Manufacturing; (4) Production and Supply of Electricity, Heat, Gas, and Water; (5) Construction; (6) Transportation, Storage, and Postal Services; (7) Information Transmission, Computer Services, and Software; (8) Wholesale and Retail Trade; (9) Accommodation and Catering Services; (10) Finance; (11) Real Estate; (12) Leasing and Business Services; (13) Scientific Research, Technical Services, and Geological Prospecting; (14) Water Conservancy, Environment, and Public Facilities Management; (15) Resident Services and Other Services; (16) Education; (17) Health, Social Security, and Social Welfare; (18) Culture, Sports, and Entertainment; (19) Public Administration, Social Organizations, and International Organizations.

the empirical strategy's ability to credibly identify gender-specific effects in the subsequent analysis.

3.2 Methods

The empirical analysis will apply the two-way fixed-effects model and 2SLS model to estimate the effect of gender income gaps on children's health. The specific human capital investment model is presented as follows:

$$Health_{it} = a + \beta IncRatio_{it} + \gamma X_{it} + P_P * t + T_t + u_i + e_{it} \qquad (1)$$

where $Health_{it}$ is the children's health status, measured by HAZ, dummy of stunting, or the number of sicknesses in a month. $IncRatio_{it}$ is the county level female-to-male income ratio assigned to the household. X_{it} is a set of control variables including the child's gender, age, dummy for urban or rural areas, year of education for both parents, height for both parents, number of children in the household, total household expense and income. T_t is the wave fixed effect, u_i is the individual fixed effect, and e_{it} is the error term. We also include the interaction of province dummies with time trends $P_P * t$, aiming to control for time-variant unobserved factors such as changes in policy or public services in a province. Standard errors are clustered at the county level.

Our preferred model is a 2SLS model. We use the Bartik instrument to fit the gender income ratio:

First stage: $IncRatio_{it} = \pi_0 + \pi_1 Bartik_{it} + \gamma X_{it} + P_P * t + T_t + u_i + \mu_{it}$

Second stage: $Health_{it} = a + \beta \widehat{IncRatio}_{it} + \gamma X_{it} + P_P * t + T_t + u_i + e_{it}$ (2)

where $\widehat{IncRatio}_{it}$ is the fitted value from the first stage. The remaining settings are consistent with specification (1). The summary statistic is shown in Table 1.

Insert Table 1

4. Results

4.1 Baseline results

The impact of income equality on children's HAZ is examined through fixed-effects and 2SLS models, using three different forms of measuring female-to-male labor income ratios, respectively. Table 2 shows the baseline results; a narrower gender income gap results in significantly better children's health. Equations (1)-(6) use a two-way fixed effects model. Comparing results (4)-(6) with (1)-(3), a higher significance level is obtained using the female-to-male ratio constructed from the demographic-based income rather than the general mean income. Column (7)-(9) use the 2SLS model, where the income ratio is fitted using the Bartik instrument. Appendix Table A1 reports the results of the first-stage regression. After incorporating the instrumental variable and additional control variables, the model explains about 58–64% of the variation in the female-to-male income ratio. Table 2 further presents the Kleibergen-Paap (K-P) F-statistics to confirm that the instrument is sufficiently strong to support reliable inference.

In column (9), the result shows that a one standard deviation (0.220) increase in the county female-to-male income ratio is associated with a 0.032-point increase in children's HAZ (at the 5% significance level), after including a full set of control variables and dummies. This corresponds to an increase of about 0.48 cm in the height of a 7-year-old child (the average height is 115.58 cm, with a standard deviation of 14.87) from a province ranked around the 10th percentile in terms of income equality to one at the median level, holding all other factors constant. From a developmental perspective, children in the sample grow approximately 6 cm per year between the ages of 5 and 14; therefore, the effect of changes in the gender income ratio is roughly equivalent to nearly one month's worth of height growth. Comparing the two models, the fixed effect regression gives a lower-bound estimate because measurement noise, child-to-parent feedback, and unobserved norms may mute the apparent benefit of gender-equal earnings. When we use a plausibly exogenous Bartik instrument to isolate genuine variation in the female relative income, the positive impact on children's height is about 30% larger.

Whether improvements in women's relative income can serve as a catalyst for reducing growth disparities between boys and girls is central to the self-reinforcing

effect of gender equality. We conduct subgroup regression analyses, with Table 3 reporting the differential health effects by children's gender. Across all model specifications, the positive health effects are consistently larger for girls than for boys. In our preferred specification, columns (5) and (6) show that a one standard deviation increase in the income ratio raises girls' HAZ scores by 0.046 points (significant at the 5% level), while boys' HAZ scores increase by 0.017 points (significant at the 10% level). This implies that, given the same change in the income ratio, a 7-year-old girl would grow approximately 0.431 cm more in height than a boy. These results indicate that narrowing the gender income gap can improve children's overall growth outcomes and, more notably, has a disproportionately positive effect on girls' growth. This contributes to closing the human capital gap between boys and girls in the next generation.

Insert Table 2 and Table 3

Family resources are often in competition when there is more than one child. Table 4 illustrates the health effects for families with different numbers of children. From Panel A, the sample consists of about 60% one-child households, and in general, the child health effect is stronger for one-child households than for two-child households. Further, in families with three or more children, the improvement in children's health associated with a narrowing of the gender income gap is less than half the magnitude observed in one-child families and is not statistically significant. Girls in one-child households experience the strongest height improvement effect, with a 0.032-point increase in HAZ relative to boys in one-child households. Although a direct comparison of height growth between boys and girls is unavailable, based on the height distribution statistics of the overall sample, this score corresponds to an advantage of 0.48 cm, representing a substantial improvement.

However, in two-child families, the positive health effect for girls is attenuated, and in families with three or more children, the effect for girls is not statistically significant. We investigate whether this phenomenon is due to girls being at a competitive disadvantage in multi-child households and conduct a further gender-based classification in the regressions reported in Panel B. Estimations (7) to

(10) contain households with more than two children, separately accounting for households with only boys, only girls, and mixed boys and girls. The results indicate that as gender equality in the labor market improves, households with only girls maintain a significant positive health effect. In households with mixed-gender children, girls are unable to sustain their advantage over boys in terms of resource allocation.

In general, the stronger health effect for girls in the overall sample is largely driven by the increased HAZ for girls in one-child families. From this perspective, China's family planning policies, which restricted the number of children during this period, contributed to the advancement of human capital among girls. However, in the absence of such restrictions on family size, the extent to which girls' health outcomes catch up with those of boys remains limited.

Insert Table 4

4.2 Heterogeneity

Traditional family values, also known as clan culture, have a substantial influence on household lifestyles. Male families are perceived to provide greater future economic support to the family and, under the influence of feudal ideology of "son preference," usually have greater access to family resources. Families more strongly influenced by clan culture also tend to limit women's decision-making power. The compilation and transmission of genealogy require the sustained efforts of clan members, and the presence of a genealogy book implies a higher level of clan cohesion. Therefore, we use the density of genealogy ownership as a measure of clan cultural intensity at the county level. Referring to the identification method of Cao et al. (2022), the 2010 CFPS household questionnaire investigated whether households kept a genealogy book, and 23.50% of households in the sample reported having one. Counties are ranked by the proportion of households with genealogy books and divided into two subgroups: areas with a high proportion of genealogy books are assigned a dummy variable defined as 1. Additionally, urban and rural areas have distinct social customs and production patterns. To account for these differences, the

analysis distinguishes between urban and rural residents using household information from the CFPS and introduces an interaction term with the income ratio to explore potential variations in the effects within these regions.

Table 5 shows that the extent to which improvements in women's relative income enhance children's growth outcomes is highly context-dependent. After introducing an urban–rural interaction term, results based on the girl sample show that, for a one standard deviation increase in the gender income ratio, rural girls experience a 0.019-point smaller gain in HAZ compared to urban girls. In contrast, there is no statistically significant urban–rural difference in the height effects for boys. It implies that rural social norms partially suppress the returns of female income gains on child nutrition. A similar pattern emerges along the dimension of clan culture. In counties characterized by strong clan cohesion—proxied by higher ancestral record ownership—the positive effect of women's income on girls' HAZ scores is reduced by 29%. Taken together, the findings show that improvements in women's relative income clearly translate into child health gains for daughters, and that patriarchal institutions rooted in rural livelihoods or clan traditions systematically dampen—but do not eliminate—these benefits.

Insert Table 5

Next, the heterogeneity is tested for labor-related characteristics. There is a significant difference in income structure between agricultural and non-agricultural households, since income from agricultural production is often not attributed to individual workers. The health effect of relative income equality is therefore weaker for agricultural households. Table 6 shows that, compare with girls from non-agricultural households, girls from agricultural households have a 0.038-point smaller improvement in HAZ. For boys, the difference in HAZ improvement between the two types of households is not statistically significant. A mother's income is considered an important source of bargaining power, with 34% of mothers in the sample reporting having no labor income, including those not participating in the labor market or only engaged in household labor. Surprisingly, column (3) shows that the absence of maternal labor income does not have a significant impact on the positive health effects for girls. A plausible interpretation is that the key mechanism

lies in household spending decisions being made based on children's expected income, rather than on which adult actually brings income into the household. As long as women's relative value in the labor market rises, mothers—regardless of their employment status—appear able to steer additional resources toward their daughters' nutrition.

Table 6 Panel B further shows the heterogeneity across regions in terms of industry. Regions with more advanced development of the tertiary sector tend to have a greater demand for labor, which means individuals are more fully able to participate in the labor market. We classify regions into high and low tertiary sector groups based on the median of tertiary sector value-added across counties, assigning a dummy of 1 to counties with a high share. It shows that regions with a higher share of the tertiary sector exhibit significantly greater positive effects on the health of both boys and girls.

The heterogeneity analyses above consistently demonstrate that in families more closely connected to the labor market, gender income equality has a greater impact on children's health equality. When parents perceive changes in the relative labor income, as well as an increase in labor opportunities for women, they are more likely to reconsider the allocation of household resources. This aligns with the hypothesis of the theoretical framework.

Insert Table 6

5. Robustness

5.1 Alternative test

In the alternative test, the explanatory variable is replaced by the binary variable "stunting," which is defined as 1 when HAZ<-2. The logit model is used to estimate the effect of gender income ratio on child stunting, with all variables following the main regression setting. The results are consistent with the main regressions, in which improvements in gender income ratio reduce the probability of stunting in children (Table 7 Panel A).

Height is an anthropometric measure, we replace it with a clinical measure, the child's sickness. Parents are asked the number of times their children had been sick in

the past month in the questionnaire. The average frequency of sickness in a month in our sample is 0.52 and is similar for both genders (0.50 for boys vs. 0.53 for girls). Panel B of Table 7 shows that a one standard deviation increase in the income ratio results in a 9.43% reduction in the number of monthly sicknesses for the girl's group, with a 6.4% reduction observed in the boy's group (although it is not statistically significant). These two alternative tests provide additional support for the main findings, suggesting that improvements in labor market gender equity are robust and positively associated with children's health, especially girl's health, across multiple dimensions.

Insert Table 7

5.2 Oster bounding approach

The fixed-effect model helps to mitigate the endogeneity problem, and in the robustness test, the Oster bounding approach (2019) is used to discuss the degree of influence on the regression results due to unobservable omitted variables. Oster (2019) calculates the bounds of the target value by adjusting coefficients and R-squared values. This method eliminates the added noise that may arise when confounding factors are included in the model by using information from R-squared. The bounding approach allows partial identification of the impact of the income ratio on children's health and constructs consistent bounds on its true value. If the identification set excludes zero, the regression results are robust, and conversely, if the identification set contains zero, the effect of health on work is stochastic.

Regarding the maximum R-squared value (R_{max}) setting, we followed the insights of Oster (2019) and used a value of R_{max}=1.3$\widetilde{R}$, where the cut-off value of 1.3 would allow at least 90% of results from randomized controlled trials to remain valid. In our robustness test, the treatment effect was positive, and the identification set is: $[\beta^{*'}(\min\{1.3\widetilde{R}, 1\}, 1), \widetilde{\beta}]$, where $\widetilde{\beta}$ is the coefficient estimated from the full controlled model, $\beta^{*'}$ is the coefficient estimated under the assumption of $\tilde{\delta}$=1 and $R_{max}=\min\{1.3\widetilde{R}, 1\}$.In addition, by setting β=0 (assuming that the female-to-male income ratio has no impact on children's HAZ), the proportional selection coefficient,

δ^0, can be obtained. This value indicates how much larger the omitted variables must be relative to the control variables for the treatment effect to be zero.

The same fixed-effect model with a full set of control variables and dummies as in the baseline regression is applied, and the results are reported in Table A2. The treatment effect in the uncontrolled and controlled regressions is shown in the first two columns. The control variables remain consistent with the main regression, and the change in R-squared can be observed as an improvement in the goodness of fit of the regression due to the observables. The last column also reports how much larger the role of the unobservable factors must be compared to the control variables for the treatment effect to become zero. The tests indicate that the significant positive effects obtained from the regression of either the general county-level gender income ratio or the demographic-based gender income ratio on children's HAZ are robust. The weight of unobservable needs to be more than 3.01 and 8.45 times as large as that of the control variables for the positive health effect to disappear from the regression.

5.3 Balance test and Rotemberg weights analysis

Following the guides of Goldsmith-Pinkham et al. (2020), the construction of the Bartik IV requires both exogeneity and assurance that the identifying variation is not driven by a small set of industry shares. To validate these conditions, we conducted a balance test and a Rotemberg weights analysis. Table A3 reports a series of balance tests to ensure that the Bartik instrument is not systematically correlated with other baseline factors (measured in 2010) that could also influence future trends in child health. For each baseline factor, we regressed in a county-level fixed effects model. None of the coefficients were statistically significant, confirming the exogeneity of the Bartik IV.

Table A4 implements a “leave-one-industry-out’’ version of the Bartik instrument in order to probe whether our baseline estimates are unduly influenced by a single sector that receives a large weight in the Rotemberg decomposition. For every industry that supplies either the top two highest male-weighted or female-weighted shock, we re-construct the employment-growth instrument after deleting that industry and re-normalizing the remaining 18 sectors; we then re-estimate the two-stage least-squares specification for overall child HAZ and separately for girls and boys. By

dropping a high-weight industry necessarily reduces the variance of the instrument, but the exercise is informative about how the estimated local average treatment effect shifts when the composition of the identifying variation changes.

When excluding the industries with the highest Rotemberg weights—construction for males, wholesale and retail trade for females, and manufacturing for both genders—the direction of the health effects in the overall, girls', and boys' samples remains consistent with the main regression results. Although standard errors generally increase, the pattern that girls experience greater positive health effects than boys continues to be observed. This confirms that no single male- or female-dominated industry is solely responsible for the positive child height response, reinforcing the robustness of the baseline findings. It also provides additional evidence that the magnitude of the effects systematically varies with the gender composition of the industries generating the income shocks.

6. Conclusion

The study offers important insights into the self-reinforcing effects of labor market equality improvements, leading to increased human capital accumulation and greater health equality among children in the next generation. The empirical results indicate that a narrower gender income gap has a significant positive impact on children's health across various dimensions. Importantly, girls seem to benefit more from these improvements, suggesting a positive intergenerational transmission of gender health equality in a favorable labor market environment.

However, the study also reveals certain disparities in health effects based on different household and cultural characteristics. Families residing in rural areas, with mothers who do not have income reporting, or who only work in agriculture, do not participate fully in the labor market and therefore have relatively weak child health effects when the labor market environment improves. The influence of clan culture slightly limits the positive health effect, potentially due to its reinforcement of "son preference" in household decision-making, which undervalues the consideration of female expected earnings. This is also reflected in mixed-sex multi-child families, where girls remain at a disadvantage relative to boys in terms of competition for family resources. The robustness tests and heterogeneity analyses provide further

credibility to the study's findings, showing that endogeneity issues are unlikely to affect the direction and significance of the observed effects. In addition, both anthropometric and clinical indicators consistently demonstrate the positive impact of gender income equity on children's health.

The channel found in this paper to promote children's human capital is more likely to be “profit-driven” that moves beyond traditional models centered on maternal altruism or current parental bargaining power, which differs from Zhao and Qu's (2024) noted preference for mothers' stronger affection for their children. Realizing higher human capital investment in children, especially girls, through higher expected female income is consistent with the mechanisms suggested by Qian (2008) and Guiso and Zaccaria (2023). We model parents as rational, forward-looking agents who allocate household resources based on the expected future returns to their children's human capital. As the opportunity cost of maintaining the feudal gender paradigm increases, people are more likely to consider abandoning gender prejudices and opting for a more equal distribution of household resources. Thus, the push for gender equality is driven by changes in the labor market.

The positive self-reinforcing effect is clear, but it is subject to limiting conditions. It can only work in more urbanized living environments and under weaker traditional clan culture. Strong “son preference” and inadequate parental participation in the modern labor market can significantly weaken the self-reinforcing effect. The overall health improvement on girls is stronger than on boys requiring a modern fertility awareness or coercive policy guidance. The one-child policy in China over the past decade has led to dominant one-child families, reducing opportunities for intrahousehold competition. When boys are present in the family, the self-reinforcing intergenerational transmission of gender equality is more likely to break down.

Our work complements research to identify factors that influence children's health to support policymakers in developing effective strategies. Some studies have shown that factors contributing to human capital accumulation, such as agricultural mechanization or minimum wage increases, have not equally benefited low-skilled female labors (Kim & Williams, 2021; Vivek Moorthy, 2022). Therefore, along with economic development, authorities should give attention to alleviating gender income inequality. Only then can they bring benefits to the health equality of the next

generation and address gender imbalances in the population. As birth policies change, allowing for more multi-child families in the future, there will be concerns about the allocation of family resources for girls. To achieve a positive reproduction effect through the profit-driven channels, there is leverage value in creating more female jobs in areas with strong gender paradigms. More diverse industries moving into less economically developed areas could be a good opportunity to provide more positions for women. We also call on policymakers to introduce more incentives to promote modern fertility concepts and to enhance social support for women, thereby improving their expected income and intra-household bargaining power.

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Self-Reinforcing Gender Equality from Labor Market to Child Health

Table and Figure

Figure

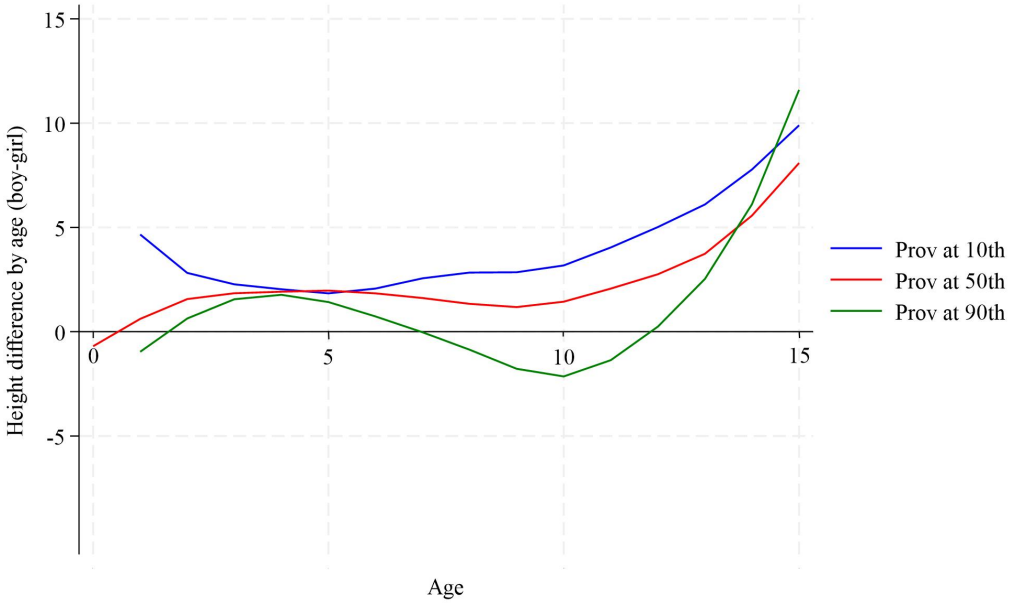

Figure 1 Gender height difference across ages by province

Note: The Y axis is the gender height difference (boy-girl) by age at the province level aggregated from CFPS. The X axis is the child age.

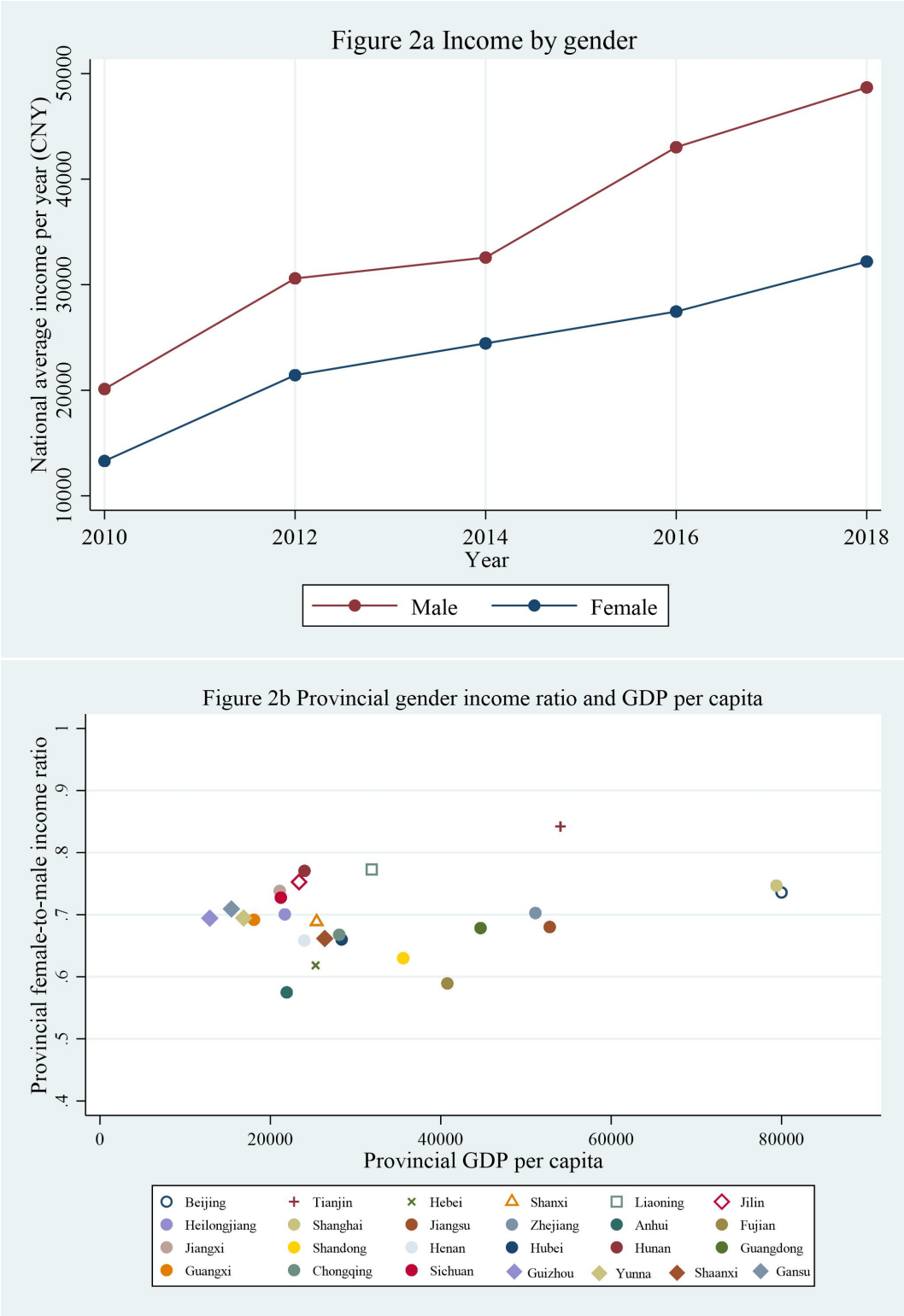

Figure 2 Female-to-male income profile 2010 to 2018

Note: Figure 2a shows the annual national average income by gender. The Y axis is the personal labor income aggregated from China Family Panel Survey CFPS (in CNY). Figure 2b shows the relationship between provincial female-to-male income ratio and GDP per capita. The Y axis is the female-to-male

income ratio at province level aggregated from CFPS. The X-axis is the average GDP per capita at the provincial level between 2010 and 2018, with data from the National Bureau of Statistics.

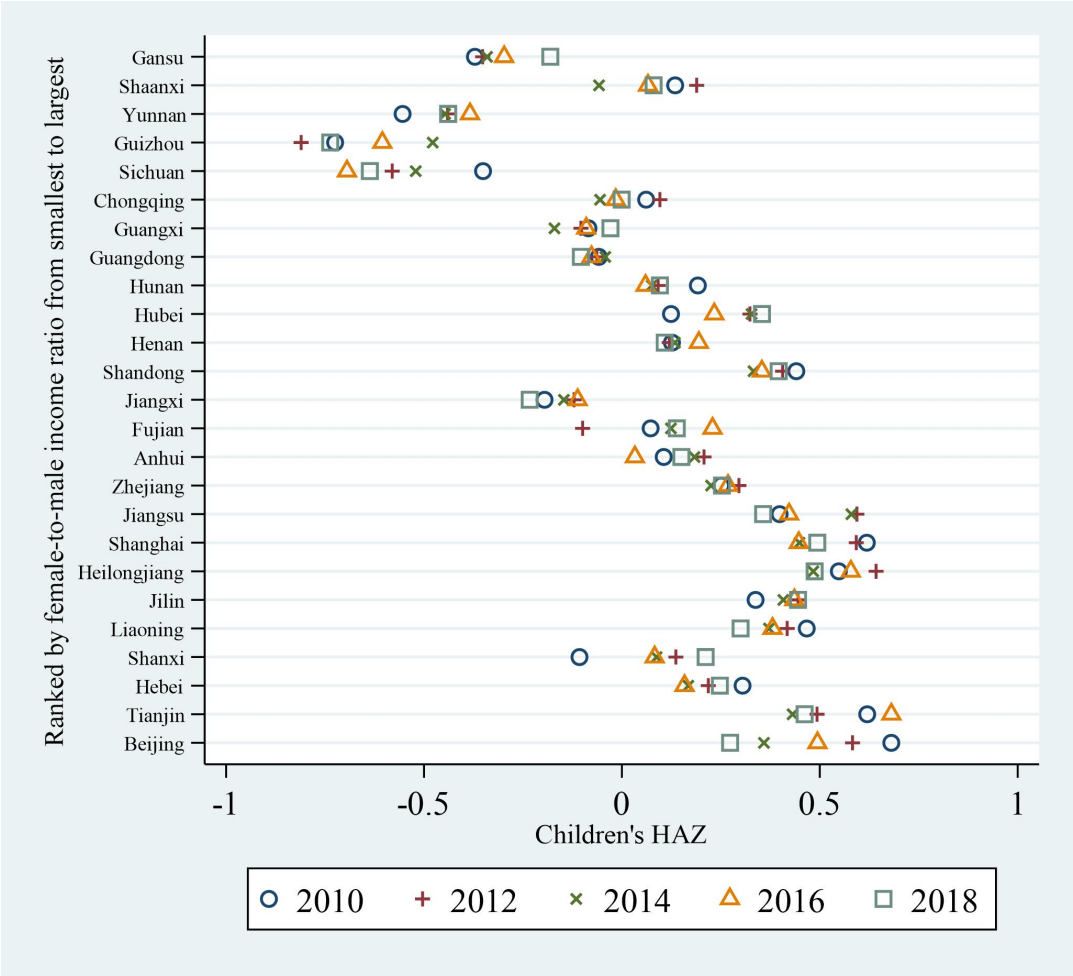

Figure 3 The distribution of HAZ by province

Note: Figure 3 shows the distribution of children's HAZ by province. Y axis is the provinces ranked by female-to-male income ratio from smallest to largest. X axis children's HAZ at province level calculated from CFPS.

Table 1 Summary statistic

Variables	Mean	Std. Dev.	Min	Max	Num of Obs
Children's height for age Z-score (HAZ)	0.000	0.999	-4.979	4.800	20,629
"Bartik instrument" generated by difference of employment growth rate	0.641	0.794	-0.912	3.581	22,956
Female-to-male income ratio generated by general county income	0.687	0.220	0.285	1.043	22,956
Female-to-male income ratio generated by demographic characteristic	0.676	0.239	0.295	1.228	22,956
Children's gender (girl=1)	0.472	0.499	0	1	22,848
Children's age	7.307	4.489	0	16	22,951
Urban and rural households (rural=1)	0.597	0.491	0	1	22,456
Households with genealogy	0.231	0.430	0	1	8,960
Number of children	1.511	0.788	1	9	22,956
Mother's years of education	7.512	4.549	0	23	18,754
Father's years of education	8.392	4.098	0	23	18,443
Mother's height (cm)	159.239	5.671	65	190	18,168
Father's height (cm)	169.780	6.116	116	202	18,019
Log of total household expense (Yuan per year)	10.210	0.846	8.002	12.221	19,138
Log of total household income (Yuan per year)	10.474	0.895	8.278	13.906	20,274
Number of children's sicknesses in the past month	0.520	0.939	0	11	17,016
Stunting in children	0.074	0.262	0	1	22,959

Source: China Family Panel Survey.

Table

Table 2 The effect of female-to-male income ratio on children's HAZ

	General county income			Income by demographic			Bartik instrument (2SLS)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Income ftm	0.110*	0.113*	0.112*	0.210***	0.148**	0.151**	0.157**	0.146**	0.146**
	(0.066)	(0.059)	(0.065)	(0.074)	(0.079)	(0.079)	(0.070)	(0.068)	(0.072)
Trend		Yes	Yes		Yes	Yes		Yes	Yes
Year dummy		Yes	Yes		Yes	Yes		Yes	Yes
Province*Trend			Yes			Yes			Yes
Num of Obs	17,735	17,735	17,729	17,693	17,693	17,687	17640	17640	17640
R-squared	0.151	0.189	0.195	0.154	0.189	0.203	0.141	0.171	0.214
K-P F statistic							23.214	16.380	70.047

Note: Each column represents the coefficient in a separate regression. Column (1)-(6) use a two-way fixed effect model. Column (7)-(9) use a 2SLS model, the first stage results are reported in Appendix Table A1. The dependent variable is the children's HAZ. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01

Table 3 The effect of female-to-male income ratio on children's HAZ by gender

	General county income		Income by demographic		Bartik instrument (2SLS)	
	Girl	Boy	Girl	Boy	Girl	Boy
	(1)	(2)	(3)	(4)	(5)	(6)
Income ftm	0.157**	0.069	0.182***	0.120***	0.210***	0.078*

	(0.070)	(0.070)	(0.050)	(0.055)	(0.056)	(0.045)
Num of Observations	8,334	9,306	8,334	9,306	8,334	9,306
R-squared	0.143	0.157	0.145	0.157	0.175	0.198
K-P F statistic					58.356	56.809

Note: All regressions in the columns include the full set of control variables and interaction terms. Column (1)-(4) use a two-way fixed effect model. Column (5)-(6) use a 2SLS model. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Table 4 The effect of female-to-male income ratio on children's HAZ by children size

	One child		Two children		Three children +	
Panel A	(1)	(2)	(3)	(4)	(5)	(6)
Income ratio	0.164***	0.091	0.142***	0.094	0.081	0.076*
	(0.053)	(0.061)	(0.044)	(0.064)	(0.061)	(0.045)
Income ratio*girl		0.146**		0.097*		0.020
		(0.070)		(0.055)		(0.048)
Num of Observations	10,168	10,168	5,739	5,739	1,733	1,733
R-square	0.186	0.184	0.165	0.167	0.182	0.130
K-P F statistic	38.021	36.727	66.301	21.070	37.564	21.085
Panel B	All boys	All girls	Boys and girls			
Family with 2+ children	(7)	(8)	(9)	(10)		
Income ratio	0.081	0.216***	0.125*	0.105		
	(0.134)	(0.083)	(0.071)	(0.073)		
Income ratio*girl				0.034		
				(0.051)		
Num of Observations	1,287	1,066	4,902	4,902		
R-square	0.219	0.178	0.190	0.172		
K-P F statistic	20.701	25.767	61.719	21.039		

Note: All regressions in the columns use 2SLS models, including the full set of control variables and interaction terms. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Table 5 The effect of female-to-male income ratio on children's HAZ by living place

	Overall	Girl	Boy
	(1)	(2)	(3)
Panel A Urban-rural			
Income ftm	0.176**	0.259***	0.096
	(0.068)	(0.076)	(0.074)
Income ftm*Rural	-0.050	-0.089*	0.024
	(0.045)	(0.047)	(0.046)
Num of Observations	17,640	8,334	9,306
R-squared	0.151	0.146	0.158
K-P F statistic	48.174	20.566	34.772
	(4)	(5)	(6)
Panel B Culture norms			
Income ftm	0.177***	0.258***	0.095
	(0.051)	(0.057)	(0.065)
Income ftm*Genealogy	-0.042	-0.075	-0.035
	(0.050)	(0.062)	(0.060)
Num of Observations	17,640	8,334	9,306
R-squared	0.175	0.174	0.174
K-P F statistic	35.826	29.462	28.930

Note: All regressions in the columns use 2SLS models, including the full set of control variables and interaction terms. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Table 6 The effect of female-to-male income ratio on children's HAZ by work conditions

	Girl	Boy	Girl	Boy
	(1)	(2)	(3)	(4)
Panel A Household work conditions				

Income ftm	0.282***	0.081*	0.218***	0.082
	(0.066)	(0.042)	(0.057)	(0.053)
Income ftm*Agriculture	-0.173***	-0.016		
	(0.049)	(0.032)		
Income ftm*Mum no income			-0.039	-0.011
			(0.030)	(0.031)
Num of Observations	8,334	9,306	8,334	9,306
R-squared	0.146	0.159	0.144	0.197
K-P F statistic	29.606	29.335	26.267	23.008
	(5)	(6)		
Panel B Regional work conditions				
Income ftm	0.088	-0.040		
	(0.077)	(0.071)		
Income ftm*Industry	0.198**	0.226**		
	(0.095)	(0.098)		
Num of Observations	8,334	9,306		
R-squared	0.153	0.162		
K-P F statistic	20.609	26.405		

Note: All regressions in the columns use 2SLS models, including the full set of control variables and interaction terms. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Table 7 Alternative test on number of sickness and stunting

	All	Girl	Boy
	(1)	(2)	(3)
Panel A Stunting			
Income ratio	-0.127**	-0.147**	-0.107*
	(0.056)	(0.059)	(0.056)
Num of Observations	15,550	7,336	8,214
R-square	0.174	0.174	0.181
K-P F statistic	85.255	78.141	64.240

Panel B Number of sicknesses			
Income ratio	-0.040***	-0.050***	-0.032
	(0.014)	(0.012)	(0.019)
Num of Observations	18,355	8,676	9,679
R-square	0.157	0.156	0.160
K-P F statistic	77.147	65.598	63.453

Note: All regressions in the columns use 2SLS models, including the full set of control variables and interaction terms. The dependent variable is the dummy of stunting in Panel A, and the number of sicknesses in Panel B. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Appendix

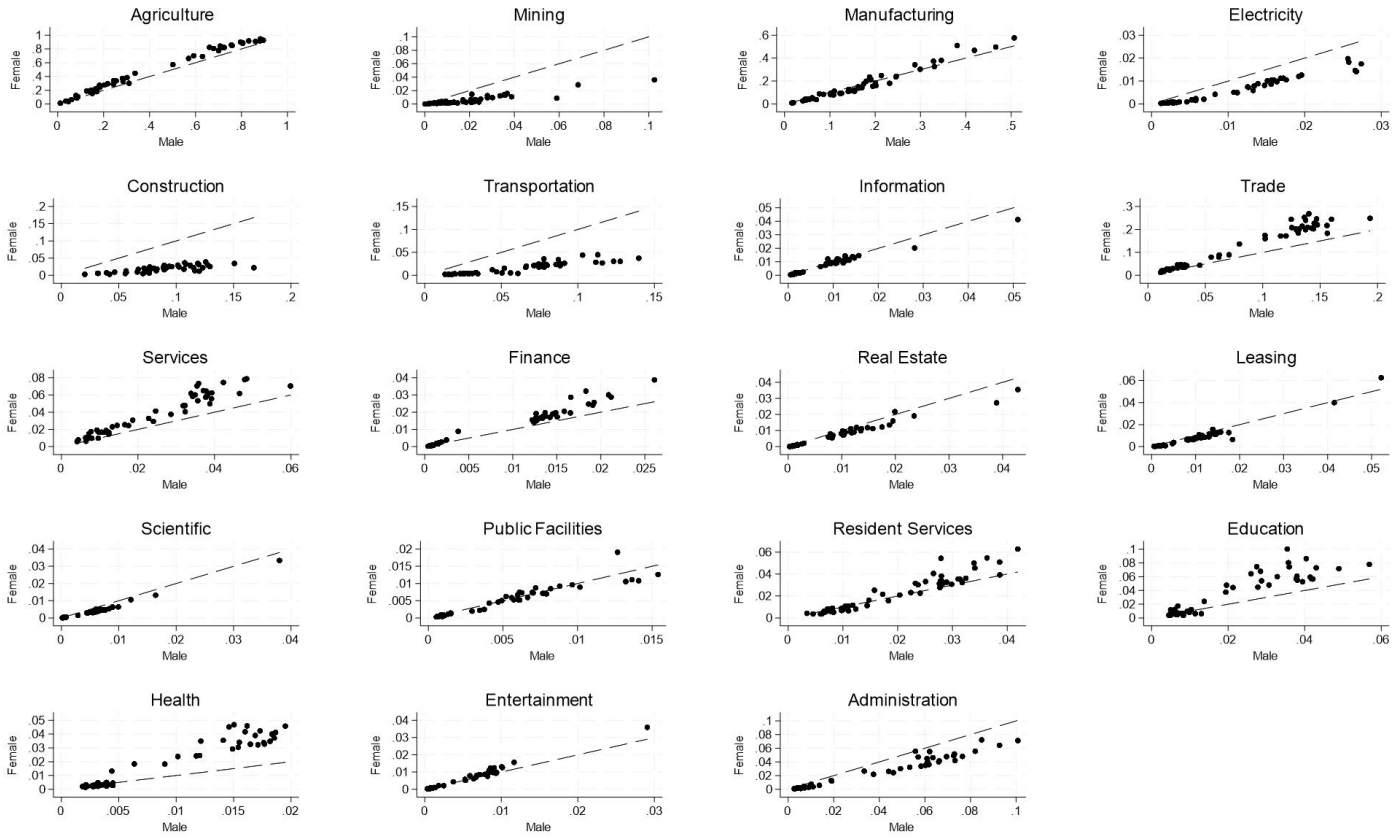

Figure A1 Gender distribution across sectors

Source: The 2010 census.

Table A1 First stage of the Bartik instrument method

	(1)	(2)	(3)
Employment growth difference	-0.428***	-0.313***	-0.334***
	(0.063)	(0.069)	(0.060)
Control variables	Yes	Yes	Yes
Trend		Yes	Yes
Year dummy		Yes	Yes
County dummy		Yes	Yes
Province*Trend			Yes

Num of Observations	18,355	18,355	18,349
R-squared	0.581	0.631	0.642

Note: Each column represents the coefficient in a separate regression. Columns (1) to (3) correspond to the first-stage results of columns (7) to (9) in Table 2, respectively. * p<0.10, ** p<0.05, *** p<0.01

Table A2 Oster bounding approach

	Uncontrolled effect β, (se), [R]	Controlled effect β, (se), [R]	Identified set	Exclude zero	δ for β=0
Dependent variable: Child HAZ					
General county income ratio	0.205*** (0.043) [0.005]	0.114** (0.075) [0.188]	[0.095, 0.114]	Yes	3.01
Income ratio by demographic	0.242*** (0.048) [0.007]	0.151** (0.071) [0.189]	[0.151, 0.163]	Yes	8.452

Note: Uncontrolled models only control female-to-male income ratio, controlled models include the full set of control variables and dummies. Results in column 3 and 5 are computed using Oster bounding approach (2019) Stata code, psacalc. * p<0.10, ** p<0.05, *** p<0.01.

Table A3 Balance test

Initial Bartik Instrument			
household level factors		county-level factors	
household expense	0.003	Education for mother	0.003
	(0.002)		(0.003)
household income	0.005	Education for father	0.003
	(0.004)		(0.003)
		Age for mother	-0.002
			(0.002)
		Age for father	-0.003
			(0.002)

Note: Each coefficient is estimated by a separate regression. All regressions include county fixed effects. Clustered standard errors at the county level are shown in the brackets. * p<0.10, ** p<0.05, *** p<0.01.

Table A4 Leave-one-industry-out test

	Industry	Rotemberg	Effect (se)		
			Overall	Girl	Boy
Largest weights in male labor	Construction	0.411	0.453*	0.692*	0.214*
			(0.266)	(0.366)	(0.128)
	Manufacturing	0.178	0.128	0.273	0.065
			(0.088)	(0.175)	(0.055)
Largest weights in female labor	Wholesale and Retail Trade	0.419	0.119*	0.172**	0.090
			(0.072)	(0.074)	(0.069)
	Manufacturing	0.240	-	-	-

Note: The Effect columns report the estimates of the effect of the reconstructed employment growth rate in which we drop the industry of each row and re-normalize the sum of the 18 industries' share of employment growth to be 1. The results using the Bartik IV excluding the manufacturing sector have already been presented earlier, so they are not repeated in the last row.