Ignorance is Whose Bliss: The Repeal of Compulsory Premarital Health Examinations and Marital Outcomes in Rural China *

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Abstract

Health matters for marital outcomes, but information about health may be hidden until marriage. Our matching model, considering socioeconomic status (SES) and health, reveals that removing health information shifts sorting towards SES, reducing child health and welfare, especially for those with low SES. Empirical evidence from China, where compulsory premarital health exams were repealed, confirms this decline in postmarital subjective well-being, primarily driven by decreased child health associated with health-based sorting. Women and low-SES individuals suffer the most, indicating persistent gender and wealth disparities.

Keywords: Premarital Health Examination; Subjective Well-being; Assortative Matching; Sorting Tradeoff; Inequality.

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1 Introduction

The health of both partners in a marriage often plays a complementary role in enhancing the production of marital surplus, with the physical well-being of prospective children being a critical component. Yet, the health status of prospective partners may remain elusive before marriage, especially in regions grappling with the burden of infectious diseases such as AIDS and viral hepatitis.[1] This situation is a likely reason why premarital health tests featuring a variety of examination items are common in many African and Middle Eastern countries (Rennie and Mupenda, 2008).[2] Despite the potential benefits, concerns have arisen regarding the potential adverse effects of disclosing health information, such as diminishing the marriage prospects of disadvantaged groups and exacerbating societal inequalities. This raises a compelling question: how does the availability or lack of health-related information impact the overall level and distribution of welfare in the marriage market? The answer to this question is not straightforward, as the effect of withholding or sharing information about one particular trait can reverberate throughout the entire marriage market featuring multiple traits through changes in matching patterns. To further complicate matters, people with different wealth or income status may differ in their accessibility to the alternative means of revealing their health status.

We attempt to address this question theoretically and empirically in the context of China. Due to the prevalence of infectious diseases, e.g., hepatitis B, China gradually adopted compulsory premarital health examinations (PHE) in the mid-1980s and expanded the coverage thereafter until 2003[3] — when the compulsory PHE was repealed. However, the ensuing

[1] According to statistics from the American Foundation for AIDS Research, 1.3 million people were newly infected with HIV in 2022, with 630,000 people succumbing to AIDS-related causes. Furthermore, new HIV infections are on the rise in regions such as Central Asia, the Middle East, and North Africa (see https://www.amfar.org/About-HIV-and-AIDS/Facts-and-Stats/Statistics--Worldwide/). Additionally, a report from the World Health Organization (WHO) states that viral hepatitis claimed 1.34 million lives in 2015, with mortality rates continuing to increase (WHO, 2017).

[2] For further details, refer to https://www.opensocietyfoundations.org/uploads/f531df52-0c99-46af-8968-8e25ab34c951/mandatory-premarital-hiv-testing-20100513.pdf.

[3] The WHO estimates that 87 million people in China (approximately 6% of the total population) are

increase in birth defects has recently rattled scholars and policy-makers, prompting calls for studies to determine how this policy-induced increase in matching uncertainty affects social welfare.[4]

We first construct a simple two-dimensional transferable utility (TU) matching model in which individuals of both genders are characterized by two traits: health and socioeconomic status (SES). Both traits exhibit the property of log complementarity to ensure assortative sorting. We consider three cases: (1) health is observable; (2) health is unobservable but high-SES people can afford to reveal their health status while those with low-SES cannot; (3) health is unobservable and no one can send the signal. By comparing these three cases, we observe a sorting tradeoff: a lack of health information generally reduces positive sorting on health but has the opposite effect on SES. Given that parental health is complementary in producing child health, the expected health of prospective children decreases due to reduced sorting by health, and the average marital welfare declines. Moreover, the distribution of welfare loss is not balanced. The welfare deterioration effect is particularly strong in healthy people because sorting on health decreases. More interestingly, the model predicts that people with low SES, on average, tend to suffer more than do high-SES people because increased sorting on SES is favorable to the high-SES individuals who also likely have alternative means of revealing their health status. The above model predictions also shed light on the differential effects on the two genders. We assume symmetric marital payoffs and population distribution across types for simplicity. If we relax this assumption by allowing the more realistic case where low-SES women proportionally outnumber low-SES men, the prediction on entrenched disparity by SES implies that women, on average, would suffer more from the lack of information. Furthermore, considering that women usually play a primary role in child care, the welfare deterioration would be even more salient for women

infected with hepatitis B virus (HBV), which can cause severe liver damage (see https://www.who.int/China/health-topics/hepatitis).

[4]For example, a Chinese news report (https://www.thepaper.cn/newsDetail_forward_1527556) states that the newborn defect rate doubled within the decade after 2003, and an academician in medical science of the Chinese Academy of Sciences, Junbo Ge, urged a change in PHE policy.

than for men if we relax the assumption of fully transferable utility.

We test model predictions by leveraging the repeal of mandatory PHE in China as a quasi-natural experiment. Our approach employs a difference-in-differences (DID) design, capitalizing on the considerable variation in PHE adoption rates among provinces prior to the repeal. Arguably, this variation is primarily attributed to differences in the accessibility of relevant medical facilities within subprovince administrative units rather than individual choices. Following the repeal, PHE rates plummeted to nearly zero in all provinces. Consequently, regions with initially high adoption rates witnessed substantial declines. Therefore, we anticipate a more pronounced impact on the marriage market in areas with higher pre-repeal PHE rates.

Our primary data source is drawn from the China Family Panel Survey (CFPS),[5] which contains important measures of marital utility—namely, subjective well-being (SWB) measures. SWB itself is an ultimate goal of economic development (Frey and Stutzer, 2002), and a growing economics literature uses self-reported SWB as a proxy for individual utility (Aghion et al., 2016; Allcott et al., 2020).[6] Our DID estimation shows that the repeal of compulsory PHE resulted in a significant decrease in self-reported SWB. Specifically, a one standard deviation increase in the pre-reform PHE ratio led to a 5.1% decrease in self-reported life satisfaction, a 4.0% decrease in self-reported family satisfaction, and a 8.2% decrease in self-reported happiness. The significant effects are supported by a battery of tests, including a parallel trend test, consideration of confounding shocks, and a permutation test.

Next, we investigate the underlying mechanisms behind our initial findings. In line with our theoretical predictions, the repeal of PHE led to a significant increase in the difference in health between spouses and a reduction in the difference in SES between spouses, suggesting decreased sorting by health and increased sorting by SES. Additionally, we observe a decline

[5]The CFPS can be considered the Chinese analogue of the Panel Study of Income Dynamics in the US.

[6]The sufficient statistics approach is another way to measure welfare gain or loss. For example, Huang et al. (2023) use the sufficient statistics approach to calculate the welfare loss caused by China's one-child policy–induced marriage distortions.

in the average health of children, which is associated with these changes in the sorting pattern.

We delve deeper into the inequality implications of the model by conducting a heterogeneity analysis of the policy's effects. We find that the decline in SWB is more pronounced among individuals with better health and lower SES. These findings align with the predictions of our model.

Overall, our findings corroborate the theoretical analysis on the sorting tradeoff. As information on health is unobservable, matching efficiency, reflected in overall SWB, likely decreases. Moreover, equity may not be achieved. The policy change, despite its aims to protect less healthy individuals, may exacerbate inequality across individuals with different SES. A particularly concerning finding is that a major channel for the decline in SWB is a decline in child health associated with decreased sorting by health. As the health of subsequent generations has important demographic and economic implications, this finding merits more attention from academia and policy-makers.

Related literature. This paper examines a novel connection between marriage and health and contributes to three strands of the literature. The first is the literature on the role of health and health tests in the marriage market. Prior studies have demonstrated the existence of positive marital sorting based on health status and behaviors (e.g., Clark and Etile, 2006; Davillas and Pudney, 2017; and Banks et al., 2021). Chiappori et al. (2020a) considered health-related behaviors and attitudes such as smoking as one trait in a multidimensional transferable utility model without information asymmetry. While studies on health tests are on the rise in economics, most are focused on individual behaviors and disease prevention (Petersen and White, 1990; Thornton, 2008; Ganczak, 2010; Delavande and Kohler, 2012; Baird et al., 2014; Wilson et al., 2014; Beegle et al., 2015; Gong, 2015; Saffi and Howard, 2015). Among the few exceptions, Buckles et al. (2011) report that the repeal of premarital blood test requirements in the US increased marriage rates, and they interpreted this change as a response to changes in the cost of marriage. Angelucci and Bennett (2021) find a high-frequency "opt-out" HIV testing intervention increased marriage and pregnancy rates and interpret it as a result of reducing information asymmetry. Our paper complements their result by focusing on a direct measure of marital surplus, i.e., SWB, in a nationwide setting. Moreover, we provide evidence for changes in sorting patterns and child health as the channels for the observed effects, which not only further corroborates the theoretical hypothesis but also yields rich implications on the distributional impacts on welfare and health.

The second strand of related literature is that on marriage matching. Following the pioneering work of Becker (1973), numerous theoretical and empirical studies examine the determinants and patterns of marriage matching. For example, using a novel twins dataset, one recent study (Li et al., 2023) finds that men with unobserved endowments related to higher earnings marry earlier and have younger and taller wives, but such women marry later and have older husbands with higher education and income. Recent developments regarding multidimensional matching enable scholars to explore the relationship between positive assortative matching (PAM) and inequality (Greenwood et al., 2014; Hryshko et al., 2017; Chiappori et al., 2020b). Our paper adds to this strand of literature by modeling and providing evidence of the sorting tradeoff between health and SES and incorporating the role of information. Becker (1981) notes that information asymmetry can be a hurdle in marriage matching, but empirical studies on this topic remain scarce (Jung and Sim, 2020).

Finally, our findings add to the evidence of the effects of public policy on SWB. SWB measures have been increasingly used as proxies for utility in evaluating tax policies (Gruber and Mullainathan, 2005), political institutions (Frey and Stutzer, 2000), the Moving to Opportunity program in the US (Ludwig et al., 2012), and public events (Dolan et al., 2019; Clark et al., 2020). SWB has also been used in the marriage market literature as a measure of the match quality and overall valuation of marriage (Weiss and Willis, 1997; Friedberg and Stern, 2014; Chiappori et al., 2018). We follow these studies and use SWB to measure the marital surplus, which lacks direct objective measures.

The remainder of this paper is organized as follows. Section 2 provides background information on the PHE. Section 3 describes the theoretical model. Sections 4 and 5 describe the data and empirical strategy, respectively. Section 6 presents the results, including baseline results and those from robustness checks. Section 7 explores underlying channels for the main findings, Section 8 tests for the inequality implications, and Section 9 concludes.

2 Background

China started to forbid marriages of individuals with uncured leprosy or venerism and of proximity of blood and introduced the PHE in the mid-1980s. The requirement of the PHE as a prerequisite for marriage registration was not explicitly stipulated until 1994 when the *Regulation on Marriage Registration* was issued.[7] This regulation stressed that marriage was forbidden for individuals with "diseases not suitable for marriage" (see Article 13), referring to the *Marriage Law of the People's Republic of China* issued in 1981 (Article 6).[8] Nonmarriageable diseases, specified later in the *Guidance for Premarital Health Examination*,[9] included severe genetic diseases, infectious diseases such as HIV, gonorrhea, leprosy and others affecting fertility, neuropathy, and diseases of vital organs.[10]

The PHE became compulsory nationwide in 1995, as mandated in the *Law on Maternal and Infant Health Care*.[11] Local governments are allowed a certain degree of flexibility in the implementation considering the variation in local medical facilities. In principle, a standard PHE checklist starts with doctor inquiries about hereditary illness and problems that might jeopardize parenting abilities, such as learning disorders and psychiatric problems, and proceeds with not only routine examinations on height, weight, and blood pressure but also fairly comprehensive blood and laboratory tests. The blood tests usually include a full blood count, liver function tests, and testing for hepatitis B surface antigen. Laboratory examinations are performed for gonococcus and sometimes trichomonads and chlamydia. Chest radiography is sometimes obtained, and abdominal ultrasonography is performed. Then,

[7] The file (in Chinese) can be found at https://www.flfgk.com/detail/9d601c9ed70b0a429612cf4a39b476f9.html.

[8] The file (in Chinese) can be found at http://www.law-lib.com/law/law_view.asp?id=44312.

[9] The guidance was issued by the Ministry of Health on July 31, 1997.

[10] The file (in Chinese) can be found at https://ncaids.chinacdc.cn/xxgx/zcwj/201804/t20180419_164348.htm.

[11] It was enacted on June 1, 1995, by the China State Council. The file (in Chinese) can be found at http://www.law-lib.com/law/law_view.asp?id=547.

couples receive hour-long premarital health instructions. Upon receiving the examination results, couples meeting the requirements are issued a certificate of health for marriage. In other instances, marriage must be postponed to allow for some form of treatment or counseling. Couples with severe psychiatric diseases or with low intelligence must agree to receive permanent contraception. Some people, even those who receive certificates, may opt not to marry after learning more about each other's health status.

Nevertheless, the requirement cannot be implemented universally, especially in rural areas. The marriage certificate is issued by the local government of the couples' household registration location, with government-designated hospitals providing the PHE.[12] Rural couples would have to go to their township government for a certificate. In townships lacking health facilities, the compulsory PHE cannot be strictly implemented. Thus, the adoption rate of the PHE varied. The average PHE rate peaked at 68% in 2002.

The PHE was unappealing to many couples partly because of its associated cost. Apart from the cost of undergoing examinations, the PHE involves nontrivial time and transportation costs.[13] Approximately 54.2% of respondents in a survey conducted in Beijing cited the associated hassle as their reason for not taking the PHE.[14] Those costs are especially salient for rural residents. Moreover, the psychological cost weighs heavily on people unwilling to have their illness revealed (Sun, 2006; Buckles et al., 2011).[15] More than half of respondents in a survey conducted in Shanghai reported that they would refuse a PHE for fear of being discovered to be sick.[16]

[12]China has implemented a strict household registration system since 1950s, which ties an individual to a small administrative unit such as a street or village. An individual's household registration status (*hukou*) follows his/her parents' and is linked with a variety of social welfare and rights. *Hukou* is difficult to change, especially for rural residents.

[13]The PHE fee differs by location. According to reports in 2006 and 2007, it usually cost a couple 200 to 400 RMB yuan (approximately 30 to 50 USD) to complete the PHE. This cost is nontrivial given that the monthly household income per capita was approximately 100 USD in 2005. Some anecdotal evidence (in Chinese) can be seen at https://www.cctv.com/law/20060221/100657.shtml.

[14]See the news report (in Chinese) at http://news.sina.com.cn/o/2005-03-05/08175272366s.shtml.

[15]See some anecdotal evidence (in Chinese) at http://www.cpwnews.com/content-26-3141-1.html.

[16]See the news report (in Chinese) at http://news.sina.com.cn/s/2004-09-30/01473805427s.shtml.

The compulsory and comprehensive nature of the PHE stirred ethical concerns. There were criticisms that the PHE weakened the protection of the basic rights of less healthy people and entrenched their disadvantage. In response, the State Council lifted the requirement of qualified medical examination results for marriage registration in a new version of the *Regulation on Marriage Registration*, which was enacted on October 1, 2003.[17] As shown in Figure 1, the PHE rate plummeted from 68% in 2002 to a mere 2.7% in 2004 and remained low until approximately 2008. An upswing in birth defects after the repeal caught the attention of the media and aroused concerns that the repeal of the compulsory PHE conflicted with the spirit of the *Law on Maternal and Infant Health Care*.[18] Some local governments started to provide incentives for new couples to take the PHE in 2008, which may explain the modest increase in the PHE rate after 2008 (Zhou et al., 2015).

The impact of the PHE repeal varied widely among provinces because the pre-2003 PHE adoption rates differed greatly. Our data show that in 2002, PHE rates exceeded 90% in Beijing and Shanghai but were below 50% in poorer provinces, such as Qinghai, Henan, and Shanxi.[19] Zhou et al. (2015) show that the average PHE rates between 1995 and 2002 in the east, middle, west, and northeast of China were 72.6%, 42.9%, 58%, and 77%, respectively. Figure 2 plots the provincial reduction in PHE rates between the 2004-2007 average and 2002 against the rates in 2002. Except for the point corresponding to Tibet, most of the points lie near the 45-degree line, showing that provinces with PHE rates one percentage point higher before 2003 show reductions of almost the same magnitude after the repeal.

3 A TU Model of the Marriage Market

We first set up a simple two-dimensional TU model to analyze the following three cases: (1) health is observable; (2) health is unobservable with costly signaling; (3) health

[17]The file (in Chinese) can be found at `https://www.gov.cn/zwgk/2005-05/23/content_167.htm`.

[18]Some evidence from the media shows that the ratio of newborn infants with defects doubled within the decade after 2003 (see the article (in Chinese) at `https://www.thepaper.cn/newsDetail_forward_1527556`).

[19]The only available data that we have on provincial-level PHE rates start in 2002.

is unobservable without signaling. Then, we analyze how the repeal of the compulsory PHE affects the equilibrium outcome by comparing the three cases.

3.1 Model Setup

We consider two populations, men and women, with women on the short side, denoting the female-to-male ratio as μ. Agents of both genders possess two fixed binary traits: health and SES. Without loss of generality, we assume that the health index takes values 0 or 1, with 0 representing low health status and 1 representing high health status; similarly, we assume that the binary SES index takes values in the set $\{p, r\}$, where p stands for low SES and r stands for high SES. An agent is thus characterized by a pair (x, X) if male and (y, Y) if female, where $x, y \in \{0, 1\}$ is the agent's health status, and $X, Y \in \{p, r\}$ $(r > p > 0)$ is their SES.

Given the setup, each gender can be classified into four types: the high-rich type $(1r)$, the high-poor type $(1p)$, the low-rich type $(0r)$, and the low-poor type $(0p)$. The measure of type $1r$ males is normalized to be 1. The measures for types $1r$, $0r$, $1p$, $0p$ men (women) are, respectively, assumed to be $1(\mu)$, $1(\mu)$, $\delta_1(\delta_1\mu)$, and $\delta_2(\delta_2\mu)$, where the female-male ratio μ is assumed, for convenience, to be the same for the four types.

Note that we allow health to be either independent of or positively correlated with SES by assuming $1 \leq \delta_1 \leq \delta_2$. Thus, there is not a greater number of rich people than poor people, and, on average, poor people are not healthier than rich people. These assumptions align with reality. We denote the proportion of healthy agents in low-SES individuals as $\tilde{\delta_1}$, $\tilde{\delta_1} = \frac{\delta_1}{\delta_1+\delta_2}$. Therefore, $\tilde{\delta_1} \leq \frac{1}{2}$, i.e., the likelihood of a low-SES agent being healthy is no greater than that of a high-SES agent.

We make the following technical assumptions:

ASSUMPTION 1. $\frac{2+\delta_1}{2+\delta_1+\delta_2} < \mu < 1$.

This assumption is adopted to help us specify the equilibrium matching and marital output allocation schemes. $\mu > \frac{2+\delta_1}{2+\delta_1+\delta_2}$ ensures that there are sufficiently numerous women; that is, the first three types of men (i.e., 1r, 1p, and 0r) combined are not enough to marry all the women.[20] The assumption of $\mu < 1$ is motivated by the empirical finding that fecund women are in relative scarcity in the marriage market, even if the sex ratio at birth is more balanced. As Siow (1998) notes, the shorter fertile period of women than that of men implies that, in monogamous societies with divorce and remarriage, fecund women are relatively scarce. Moreover, most developing countries suffer from the "missing women" problem, China being a notorious example. Technically, our main results still hold if this assumption is relaxed.

In any married couple, the sum of individual utilities, determined by the partners' characteristics, consists of two parts — nonhealth marital output and child health status — the latter being determined by the couples' health. We assume that the surplus Σ generated by a match between Mr. (x, X) and Ms. (y, Y) has the form

$$\Sigma((x, X), (y, Y)) = q(x, y)t(X, Y)$$

where $q(\cdot, \cdot)$ is the production function of prospective child health status and $t(\cdot, \cdot)$ is the nonchild marital output produced from the SES of the couple. Thus, the utility functions of Mr. (x, X) and Ms. (y, Y), $u_{xX,yY}$, $v_{xX,yY}$ satisfy

$$u_{xX,yY} + v_{xX,yY} = \Sigma((x, X), (y, Y))$$

Following Chiappori et al.(2012), we adopt a multiplicative separable form of the marital output function that ensures assortative matching for type $1r$ agents. In our notation, a single person is written as being matched with ϕ, the empty set. In the absence of out-of-wedlock birth or adoption, a single man (x, X) or a single woman (y, Y) derives utility $t(X, \phi_Y)$ or $t(\phi_X, Y)$, respectively, which are normalized to be 0 following the custom in the literature.

ASSUMPTION 2. (Log Complementarity in SES and health) For any $X, Y \in \{r, p, \phi\}$, $t(X, Y)$ is increasing in both X and Y, and *log* $t(X, Y)$ is super modular, i.e., $t(r, r)t(p, p) \geq$

[20] We show in Part A of the Online Appendix that single persons must include some low-poor men. Given the measures of each type, the total measure of women is $\mu(2 + \delta_1 + \delta_2)$, which is greater than the measure of the first three types of men $1 + \delta_1 + 1$. This condition means $\mu > \frac{2+\delta_1}{2+\delta_1+\delta_2}$.

$t(r,p)t(p,r)$. Similarly, for any $x, y \in \{0,1\}$, $q(x,y)$ is increasing in both x and y. For any $x, x', y, y' \in \{0,1\}$ with $x > x', y > y'$, we assume that $q(x,y)q(x',y') \geq q(x,y')q(x',y)$.

It is conventional to assume that both men and women can be ranked by SES, which is complementary (or supermodular) in producing marital output. The complementarity property can be rationalized by increasing returns in the household production function or household public goods.[21] This assumption of complementarity in health is motivated by existing evidence. In Appendix Tables B1a and B1b, we provide evidence of the complementary role of spousal health in child health production.

We follow Smith (2006) in assuming log complementarity in $t(X,Y)$ and $q(x,y)$, which is stronger than complementarity. We make this assumption because the complementarity in SES must be sufficient such that not all the rich types would accept the poor types for good health, and similarly for health. For the ease of calculation, we assume without loss of generality that $q(x,y) = \lambda^{1-x}\lambda^{1-y},\ 0 < \lambda < 1$. That is, child health $q(.,.)$ takes values of 1, λ, λ^2, respectively, in cases with both parents being healthy, one and only one parent being healthy, and no parents being healthy.

Under assumption 2, it is possible that all rich types would accept only rich types regardless of health. This case of strict sorting by SES is not of the main interest in our setting since we investigate the impact of health information on marriage matching. For the purpose of this study, we focus on the case where health matters to the extent that the assortative matching is not strictly by SES, which is guaranteed by the following assumption.

ASSUMPTION 3. $t(r,p) - t(p,p) > \lambda[t(r,r) - t(r,p)]$.

That is, the marginal contribution of a high-SES spouse to a marriage with a type $1p$ agent is greater than the marginal contribution of a high-SES spouse to a marriage with a type $0r$ agent, conditional on the same health status. This assumption ensures that the marital surplus between type $1r$ and type $1p$ is greater than that between type $1r$ and type $0r$.

[21] For example, Weiss, 1997; Lam, 1988; Iyigun and Walsh, 2007; Chiappori et al., 2009; Banerjee et al., 2013; Weiss et al., 2018.

3.2 Equilibria

DEFINITION 1. Suppose a man (x, X) receives expected payoff Eu^*_{xX} and a woman (y, Y) receives payoff Ev^*_{yY} from a match and that $F^*_{xX,yY}$ ($G^*_{xX,yY}$) is the measure of type xX men (type yY women) who are married to type yY women (type xX men). A marriage market equilibrium (core) arises if the following three conditions are satisfied for all xX and yY:

1. $Eu^*_{xX} \geq t(X, \phi)$ and $Ev^*_{yY} \geq t(\phi, Y)$.

2. $Eu^*_{xX} + Ev^*_{yY} \geq Eu_{xX,yY} + Ev_{xX,yY}$ for matches between any xX and yY.

3. $F^*_{xX,yY} = G^*_{xX,yY}$, for any $xX, yY \in \{1r, 1p, 0r, 0p\}$.

The first condition requires that no individual can be worse off than they would be from remaining single. The second condition requires that the equilibrium cannot be blocked by a deviating individual (couple); that is, the equilibrium must be stable. The third condition requires that the measure of married men and women be equal in each marriage category, implying that the market clears.

3.2.1 Equilibrium if Health is Observable

Given assumptions 1-3, if both health and SES are observable, the equilibrium assignment is as follows.

$$\mathbf{A} = \left[\begin{array}{c|ccccc} & 1r & 1p & 0r & 0p & \phi \\ \hline 1r & \mu & 1-\mu & & & \\ 1p & & \delta_1\mu + \mu - 1 & (1+\delta_1)(1-\mu) & & \\ 0r & & & \mu - (1+\delta_1)(1-\mu) & (2+\delta_1)(1-\mu) & \\ 0p & & & & \delta_2\mu - (2+\delta_1)(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \end{array}\right]$$

where the element of matrix $A_{i,j}$ equals the number of type-i men who are matched with type-j women, $i, j \in \{1r, 1p, 0r, 0p\}$. The proof is in Appendix.

This case exhibits predominantly assortative matching by health and assortative matching by SES within a health status. Mixed marriages across health arise mainly from the unbalanced sex ratio in each type. The measure of mixed marriages across health is therefore

$$A(mixed\ health) = A_{1p,0r} = (1+\delta_1)(1-\mu)$$

The measure of mixed marriages across wealth status is thus

$$A(mixed\ wealth) = A_{1r,1p} + A_{1p,0r} + A_{0r,0p} = 2(2+\delta_1)(1-\mu)$$

3.2.2 Equilibrium if Health is Unobservable but High-SES Individuals Have Means to Reveal Health

If health is hidden, type 1 agents would have incentives to reveal their health status. By contrast, low-SES individuals are more likely to be deterred by the cost or have less knowledge about the available examinations. Therefore, we consider a case where only high-SES agents are able to reveal their health status.

It is easy to show that when the fraction of healthy agents among low-SES individuals is not excessively large ($\tilde{\delta_1} \leq \frac{t(r,r)-t(r,p)}{t(r,p)} \frac{\lambda}{1-\lambda}$), we have $\lambda t(r,r) > \tilde{\lambda} t(r,p)$, where $\tilde{\lambda} = \tilde{\delta_1} + (1-\tilde{\delta_1})\lambda$ is the expected child health for a type $1r$ agent who marries a type p spouse. That is, for a marriage with type $1r$ agents, the marginal contribution of a type $0r$ spouse would be greater than that of a random type p spouse. In this case, a stable matching exists, and the equilibrium assignment is as follows.

$$\mathbf{A} = \left[\begin{array}{c|cccc} & 1r & 0r & p & \phi \\ \hline 1r & \mu & 1-\mu & & \\ 0r & & 2\mu-1 & 2(1-\mu) & \\ p & & & (\delta_1+\delta_2)\mu - 2(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \end{array}\right]$$

where the element of matrix $A_{i,j}$ equals to the size of type-i men who are matched with type-j women, $i,j \in \{1r, 1p, 0r, 0p\}$. The proof is in the Appendix.

Since the health status of low-SES agents is unobservable, matches with low-SES agents would be randomized across type $1p$ and type $0p$ agents. The measure of mixed marriages across health status is

$$\begin{aligned} A(\textit{mixed health}) &= A_{1r,0r} + \tilde{\delta_1} A_{0r,p} + 2\tilde{\delta_1}(1-\tilde{\delta_1})A_{p,p} \\ &= (1-\mu)(1-2\tilde{\delta_1}) + 2\tilde{\delta_1}\delta_2\mu \end{aligned}$$

The measure of mixed marriages across wealth status is

$$A(\textit{mixed wealth}) = A_{0r,p} = 2(1-\mu)$$

Note that the condition $\tilde{\delta_1} \leq \frac{t(r,r)-t(r,p)}{t(r,p)} \frac{\lambda}{1-\lambda}$ would be violated if the utility loss from having an unhealthy child is too large (i.e., λ is too small). If the condition is violated, i.e., the fraction of healthy agents among low-SES individuals is sufficiently large relative to the utility loss from having an unhealthy child so that type r individuals would be better off by marrying a random low-SES agent than marrying a type $0r$ agent, there is no stable matching. Therefore, we do not consider this case.

3.2.3 Equilibrium if Health is Unobservable and No One Signals

If no one is able to reveal their health status, the matching problem is reduced to a one-dimensional problem. Assortative matching would arise, and matching by health would be random within an SES category. The equilibrium assignment is as follows.

$$\mathbf{A} = \left[\begin{array}{c|ccc} & r & p & \phi \\ \hline r & 2\mu & 2(1-\mu) & \\ p & & (\delta_1+\delta_2)\mu - 2(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \end{array} \right]$$

where the element of matrix $A_{i,j}$ equals the size of type-i men who are matched with type-j women, $i, j \in \{1r, 1p, 0r, 0p\}$. The proof is in the Appendix.

The measure of mixed marriages across health status is

$$A(\textit{mixed health}) = 1 + [(\delta_1+\delta_2)\mu - 2(1-\mu)]2\tilde{\delta_1}\tilde{\delta_2}$$

The measure of mixed marriages across wealth status is

$$A(\textit{mixed wealth}) = A_{0r,p} = 2(1-\mu)$$

3.3 The Role of Information

We examine the role of information by comparing the matching pattern and welfare of different groups in the three cases. Our Proposition 1 compares how the equilibrium matching pattern depends on information.

PROPOSITION 1. *Positive sorting based on SES would increase and positive sorting on health would decrease when the information on health is hidden.*

Proof. See the proof in Appendix A.2. □

This proposition is intuitive. It is easy to show that the measure of mixed marriages across health increases from case 1 to case 3 and that the measure of mixed marriages across SES decreases from case 1 to case 3.

PROPOSITION 2. *The expected child health decreases as positive sorting on health decreases.*

Proof. See the proof in Appendix A.2. □

Proposition 2 arises from the complementarity of parental health in producing child health. As sorting by health decreases, mixed-health marriages increase in number while same-health marriages decrease, and the complementarity of parental health in producing child health dictates that the average expected child health would decrease.

Associated with the decrease in average child health, the overall expected utility of married couples decreases. Intuitively, as assortative matching on health decreases, the complementarity of child health production cannot be fully realized. Even though assortative matching on SES increases, the overall matching efficiency decreases. This intuitive result can be formally stated in the following proposition.

PROPOSITION 3. *As positive sorting on health decreases, the expected equilibrium of marital output decreases.*

Proof. See the proof in Appendix A.2. □

While overall expected utility decreases, the impacts are asymmetric across types. We have the following proposition on welfare inequality.

PROPOSITION 4. *As positive sorting on health decreases,*

1. *the expected utility gap between the healthy type and unhealthy type shrinks.*
2. *the expected utility gap between high-SES and low-SES individuals likely widens unless health information is unobservable for all groups.*

Proof. See the proof in Appendix A.2. □

Proposition 4(1) is straightforward. As health is unobservable, individuals with better health lose their advantage in the marriage market: their chance of marrying a healthy spouse decreases. The expected utility of the healthy group decreases by a larger amount than that of their counterparts with worse health because of the complementarity in spousal health. Therefore, the expected utility gap across health groups shrinks.

The prediction in Proposition 4(2) arises because the high-SES group is more likely to overcome the barrier and reveal their health status. The high-health high-SES group can still reap the gains from assortative matching by health, while high-health low-SES individuals' chances of marrying a high-SES or high-health spouse decrease. Although low-health low-SES agents gain from the information "blurriness", overall, the low-SES agents suffer more. The utility gap between the rich and poor is likely entrenched. If health information is unobservable for all groups, the matching is essentially one-dimensional. High-SES agents may gain from increased sorting on SES but lose from reduced sorting on health. Their loss is greater than that of the low-SES group because the marginal effect is larger given our assumptions. Therefore, the utility gap between the rich and poor may decrease, even though both groups suffer.

Remarks on gender differences. We assume fully flexible transfers within couples and symmetric gender roles in producing child health and monetary marital output. Under these assumptions, the change in the utility gap between males and females is not straightforward but depends particularly on the population distribution among the different types. Nevertheless, the predictions above can shed some light on the utility gap between males and females. First, a possible extension of our model is to allow the within-couple transfer not to be fully flexible because females are usually the main caregiver. If this case, the predicted decrease in average child health would be more likely to adversely impact women than men. Second, our model predicts that low-SES people would suffer from more adverse effects under certain circumstances. If we relax the assumption of a symmetric population distribution between the two genders across types to allow women with low SES backgrounds to outnumber men in proportion, Proposition 4(2) implies that women, on average, would likely suffer more from the lack of health information.

3.4 Linking Model Predictions to the Empirical Setting

Now we bring the model to the empirical setting. In our empirical context, the presence of the compulsory PHE is viewed as case 1 in our model. In the scenario without the compulsory PHE, either case 2 or case 3 occurs depending on whether high-SES individuals are able to reveal their health status. In provinces with higher implementation rates of compulsory PHE, a greater number of local marriage markets are in case 1. The repeal of compulsory PHE shifts instances of case 1 to cases 2 or 3. Given Propositions 1-4 and remarks on gender differences, we have the following hypotheses.

Hypothesis 1. *Married individuals, on average, experience a larger decrease in SWB after the repeal of the compulsory PHE in areas with more prevalent PHE use ex ante than those in areas where the PHE was less common.*

We also test channels through which utility decreases. Propositions 1 and 2 lead to the following corresponding hypotheses.

Hypothesis 2. *Child health, on average, tends to decline more after the removal of the PHE in areas with a higher ex ante PHE prevalence than in areas where the PHE was less prevalent.*

Hypothesis 3. *Areas with high prior prevalence of the PHE see more marital sorting on SES and less sorting on health after the removal of the PHE than do areas where the PHE was less prevalent.*

Proposition 4 generates the following predictions on inequality.

Hypothesis 4. *The decrease in subjective well-being is larger for relatively healthy individuals than for relatively unhealthy individuals, for the poor than for the rich, and for women than for men.*

4 Data

To test the aforementioned hypotheses, we combine individual- with provincial-level data from China.

We obtained individual-level data from the China Family Panel Survey (CFPS), which is a biennial panel survey initiated in 2010 by the Institute of Social Science Survey at Peking University. The CFPS employs an implicitly stratified multistage sampling method and covers 25 of 31 provinces in China.[22] This dataset provides comprehensive information on individual economic and noneconomic conditions. The baseline survey comprised 33,600 adults and 8,990 children from 14,960 families. The follow-up surveys not only tracked the original respondents but also incorporated newly formed families into the sample.[23] Our primary dataset is from the 2014 wave because it contains more extensive information regarding individual SWB compared to other waves. Additionally, we utilize data from the 2010 and 2012 waves to impute missing data on time-invariant individual characteristics in the 2014 wave.

We focus our main analysis on the rural sample because the PHE likely plays a more important role in rural areas, where infectious diseases and birth defects are more prevalent

[22]Notably, Ningxia, Inner Mongolia, Qinghai, Hainan, Xinjiang, and Tibet are not included in the survey. Since Tibet is not part of the sample, it is not considered an outlier, as illustrated in Figure 2.

[23]For further information about the CFPS, please visit `http://www.isss.pku.edu.cn/cfps/en/index.htm`.

than in urban areas.[24]

We define rural residents as those whose *hukou* (household registration) status is rural. We keep individuals who registered for marriage between 1995 and 2007, including those who divorced during this period. As the repeal of the compulsory PHE occurred in October 2003, it is difficult to determine whether couples who married in 2003 were affected. Thus, we exclude these couples from our sample. Given that we have no information on provinces where individuals registered for marriage, we keep only individuals who had never migrated across provinces after age 12.We end up with 3,213 individuals as our main analysis sample. We further construct an urban sample based on the same criterion to use in our robustness checks.

Our primary outcome variable is SWB. SWB has been used in the marriage market literature as a measure of the matching quality and the overall valuation of marriage (Weiss and Willis, 1997; Friedberg and Stern, 2014; Chiappori et al., 2018). The CFPS contains information that can be used to construct three measures of SWB. The first is *life satisfaction*, constructed from the question "How satisfied are you with your life?" The answer has a value ranging from 1 (very unsatisfied) to 5 (very satisfied). The second is *satisfaction with family*, constructed from the survey question "How satisfied are you with your family?" This answer also takes a value from 1 (very unsatisfied) to 5 (very satisfied). The third is *happiness*, constructed from the question "How happy do you think you are?" The self-rating scores range from 0 (lowest level of happiness) to 10 (highest level of happiness). We directly use the answers to the three questions to construct the SWB measures.[25]

[24] Wang et al. (2019) document that the prevalence of HBV is 3.29% and 5.86% among urban and rural people, respectively. Furthermore, data from National Disease Surveillance Points show that the average rate of birth malformation was higher in rural areas (1.24%) than in urban areas (0.79%) from 1991 to 1995. In addition, the Report on China Birth Defect Prevention (2012) shows that the infant mortality rate due to birth defects was 4.3‰ in rural areas but 3‰ in urban areas in 2000. In the same year, maternal mortality was 0.7‰ in rural areas and 0.3‰ in urban areas, and the mortality rate of children under age 5 was 45.7‰ in rural areas and 13.8‰ in urban areas (*China Health Statistic Yearbook*, 2001).

[25] One concern is that self-reported SWB scores are not comparable across individuals, as self-reported measures are affected by other idiosyncratic factors such as the interpretation of the response categories

We gather provincial-level information from various sources. First, the PHE rate, defined as the share of PHE takers among newly married individuals in 2002, is taken from the *China Health Statistical Yearbook* (2003). We use the PHE rate to measure preabolition exposure to PHE in each province. We collect other provincial-level economic and demographic characteristics in 2002 from the *China Statistical Yearbook* (2003) and *China Health Statistical Yearbook* (2003), including GDP per capita, consumption per capita, share of the primary industry in GDP, number of health institutes per 10,000 people, number of doctors per 10,000 people, number of civil affair staff per 10,000 people, government health expenditures per capita, and infection rate of Categories A and B infectious diseases.[26]

We present the summary statistics of major variables in Table 1, in which Panel A contains the details for individual-level variables and Panel B those for province-level variables. From Panel A, we can see that the average value of *life satisfaction* is 3.768 (out of 5), the average value of *family satisfaction* is 3.864 (out of 5), and the average value of *happiness* is 7.449 (out of 10). The average PHE rate is 0.689, as shown in Panel B. Due to space limitations, we do not describe the summary statistics of other variables, the details of which are provided in Table 1.

5 Empirical Strategy

We exploit the DID strategy to estimate the causal effect of the repeal of the compulsory PHE on SWB. We combine two dimensions of variation. The first is the change between the pre- and post-PHE repeal periods. Specifically, we compare individuals who married before 2003 with those who married after 2003 (we drop individuals who married in 2003 to avoid the mixture of effects). The second source of variation comes from the different provincial PHE rates measured in 2002, one year before the policy change. The higher the PHE rate in

(Kahneman and Krueger, 2006). However, as Kahneman (1999) notes, this issue is less problematic in a practical sense when we compare two groups consisting of many individuals, in which case the reporting bias is likely to be trivial (Di Tella and MacCulloch, 2005; Kahneman and Krueger, 2006).

[26]According to China's Law on the Prevention and Control of Infectious Diseases, Categories A and B infectious diseases include AIDS, viral hepatitis, syphilis, and gonorrhea.

2002, the greater is the reduction in the precision of information cues and, hence, the more salient the impact is expected to be. To this end, we estimate the following equation:

$$y_{icpt} = \beta \times PheRate_{p,2002} \times Post2003_t + \psi_{p,2002} \times Post2003_t + \delta_c + \phi_t + \epsilon_{icpt}. \quad (1)$$

In Equation (2), i denotes the individual, c the county, p the province, and t the year of marriage. y_{icpt} is a set of outcome variables, including *life satisfaction*, *satisfaction with family*, and *happiness*. $PheRate_{p,2002}$ is the PHE rate in 2002 in province p. $Post2003_t$ is a dummy variable, where a value of 1 indicates individuals who married after 2003 and 0 those who married before 2003. δ_c and ϕ_t are county and marriage-year fixed effects, respectively, with which we control for county-specific time-invariant variables and time-specific variables. We do not separately include $PheRate_{p,2002}$ and $Post2003_t$ because they are absorbed by δ_c and ϕ_t, respectively. Note that the marriage-year fixed effects also absorb the marriage-duration fixed effects since our sample is cross-sectional data from 2014. $\psi_{p,2002}$ is a set of province-level variables measured in 2002. As $\psi_{p,2002}$ is time-invariant, we include $\psi_{p,2002} \times Post2003_t$ in the regression. The rationale for doing so is discussed in detail below. β is the parameter of greatest interest and captures the DID effects of the repeal of the PHE requirement on SWB. To handle heteroskedasticity and serial correlation within provinces and across provinces within a marriage year, we calculate the standard errors by two-way clustering over province and marriage year. We calculate standard errors using different clustering methods as a robustness check (see Section 6.2).

There are several concerns about the validity of our DID strategy. First, the PHE rate in 2002 was not randomly distributed across provinces. Thus, it could have been correlated with other provincial variables in 2002. If these variables drove changes in our outcome variables between the pre- and post-PHE repeal periods, our estimates could be biased. We investigate how the PHE rate in 2002 is correlated with other province-level variables measured in the same year. Specifically, we regress the PHE rate on other province-level variables, including GDP per capita (in logarithm), consumption per capita (in logarithm), the share of the primary industry in GDP, number of health institutes per 10,000 people, number of doctors per 10,000 people, number of civil affair staff per 10,000 people, government expenditures on health per capita (in logarithm), and infection rate of Categories A and B infectious diseases. The results are shown in Appendix Table B2. We first regress the PHE rate on these province-level variables one by one in Columns (1) to (8); the PHE rate is significantly correlated with five of eight variables. Then, we regress the PHE rate on these variables together in Column (9). Although the coefficient of each single variable is not significant, the p value of the joint F test is less than 1%. This exercise suggests that the PHE rate is indeed not randomly distributed across provinces. To address this issue, we control for the aforementioned province-level variables measured in 2002 in Equation (2), $\psi_{p,2002}$ (interacted with $Post2003_t$).

Second, the assumption on which any valid DID is predicated is that there would have been no differential time trends in the outcome variables for provinces with different PHE rates in 2002 in the absence of the policy change. We test whether this assumption holds in Section 6.2.

The third concern lies in the selection issue. The PHE repeal might have induced individuals who would not have married before the policy change to marry. If these individuals initially tended to have higher (lower) SWB, our estimates would be downward (upward) biased. To address this issue, we first investigate the impact of the PHE repeal on the provincial marriage rate. The results in Table 2 indicate that the repeal of the PHE did not significantly affect the marriage rate. Then, in the spirit of the balance test used in regression discontinuity designs (Lee and Lemieux, 2010), we check whether individual-level predetermined characteristics show systematic differences between the pre- and postreform periods across provinces. Specifically, we estimate Equation (2) but replace the outcome variable with individuals' predetermined characteristics, including the individual's parents' education and own education, the number of siblings, a minority dummy, and the logarithm of the number of weeks not living with parents before age 12. The results are shown in Table 3. No coefficients are significant, providing evidence that individuals' predetermined characteristics showed no systematic differences before and after the policy change and among provinces. [27]

[27] To supplement these results, we use the 2005 population minicensus data to conduct a similar test.

Fourth, some concurrent events occurring in the same period could be correlated with the PHE and affect the change in SWB. Three events—China's WTO entry in 2001, a college enrollment expansion after 1999 and the SARS outbreak in 2003—are notable. In Section 6.2, we show that our estimates are not driven by these three events.

In addition, we conduct other robustness checks, such as permutation tests, using different specifications and different samples to reinforce our main findings (see Section 6.2).

6 Main Results

6.1 Baseline Results: Deteriorating SWB

In Table 4, we report the baseline results. In all columns, we control for county and year-of-marriage fixed effects. The dependent variables are *life satisfaction* in Columns (1) and (2), *family satisfaction* in Columns (3) and (4), and *happiness* in Columns (5) and (6). These variables are defined in Section 4.

In the odd columns, we do not control for the interactions of provincial predetermined variables with the post dummy ($\psi_{p,2002} \times Post2003_t$). The coefficient of $PheRate_{p,2002} \times Post2003_t$ is statistically significant at the 1% level for all three outcome variables. We then add $\psi_{p,2002} \times Post2003_t$ to the regression and present the results in the even columns. The coefficient of $PheRate_{p,2002} \times Post2003_t$ is still negative for all outcome variables: equal to 1.076 (*life satisfaction*, Column 2), 0.849 (*family satisfaction*, Column 4), and 3.412 (*happiness*, Column 6). These coefficients are statistically significant. Importantly, compared with the results shown in odd columns, those in even columns are similar, thereby alleviating the concern that the PHE rate before the policy might be correlated with other provincial social and economic variables that could affect the change in individuals' SWB.

The 2005 population minicensus was conducted by China's National Bureau of Statistics and covered 1% of the total population. We have access to a 20% subsample of these survey data. The results in Appendix Table B3 show that individual characteristics are balanced in the larger sample as well. Owing to the data limitations, we can check only an individual's own education, number of siblings, and minority status when the population census data are used.

Regarding the magnitude, in comparison with the average values of the outcome variables (approximately 3.768 for *life satisfaction*, 3.864 for *family satisfaction*, and 7.449 for *happiness*), a one-standard-deviation increase (approximately 0.18) in the PHE ratio in 2002 results in a 5.1% decrease in *life satisfaction* (the product of 0.18 and -1.076, divided by 3.768), a 4.0% decrease in *family satisfaction* (the product of 0.18 and -0.849, divided by 3.864), and an 8.2% decrease in *happiness* (the product of 0.18 and -3.412, divided by 7.449).

In summary, the repeal of the compulsory PHE significantly lowered the SWB of affected individuals. This finding is consistent with Hypothesis 1 that individuals experienced a larger decrease in mental utility after the removal of the PHE in areas with higher use of the PHE before the policy change.

6.2 Robustness Checks

In this section, we conduct a battery of robustness checks to justify our main findings.

Existence of Parallel Trends. An assumption that must hold for the DID strategy to be valid is that the time trends of the outcome variables are the same across provinces in the absence of the policy change. To test this assumption, we estimate the following equation:

$$y_{icpt} = \sum_{t=1995}^{2007} \beta_t \times PheRate_{p,2002} \times Year_t + \psi_{p,2002} \times Post2003_t + \delta_c + \phi_t + \epsilon_{icpt} \tag{2}$$

where $Year_t$ is a dummy variable equal to 1 for marriage year t. The other variables remain the same as in Equation (2). We set the cohort married in 2002 as the benchmark; therefore, β_t captures the difference in outcomes in marriage cohort t in comparison with those of marriage cohort 2002 across provinces. Figure 3 plots the estimation results for Equation (3), with points representing the estimated coefficients and vertical dashed bars representing the 95% confidence intervals. For all three outcome variables, all coefficients of the interaction term $PheRate_{p,2002} \times Year_t$ before 2002 are statistically insignificant and small in magnitude, without a clear trend. The results provide evidence that parallel trends exist, thereby validating our DID assumption.

Confounding Events. Another condition for the validity of DID is that no other concurrent events are correlated with the PHE in 2002 and affect the outcome variables at the same time. Three events occurring during this period, i.e., China's entry into the WTO in 2001, the college admissions expansion after 1999 and the SARS outbreak in 2002, stand out. Furthermore, some other time-varying shocks beyond the province level may have affected the PHE and the outcome variables simultaneously. We investigate whether these events drive our main results.

First, after China entered the WTO in 2001, the domestic market faced more competition from the inflow of foreign products because of reduced import tariffs (Brandt et al., 2017). Provinces with lower PHE rates in 2002 could be less economically developed and therefore more likely to be impacted by an increase in foreign competition. If more competition reduces individuals' SWB, China's entry into the WTO could lead to downward bias in our estimates. Second, starting in 1999, China expanded college enrollment; therefore, the number of college graduates increased dramatically in 2002 (college takes at least three years), and the increase in college graduates in the labor market might have driven down wages (Li et al., 2017). Moreover, college admissions could expand more in provinces with higher PHE rates. If lower wages are linked to a lower SWB, then our estimates would be upward biased. Finally, SARS broke out in China in 2002, and provinces with higher PHE rates may have suffered less. If SARS lowered individuals' SWB, then our estimates could be underestimated.

To determine whether these three events drive our main results, we include in the regression the logarithm of total trade value (i.e., exports plus imports), the logarithm of the number of college graduates, and the logarithm of the number of SARS-infected persons[28] of each province in 2002 (interacted with a post-2003 dummy).[29] The estimation results are reported in Panel A of Table 5. The main results remain robust.

Furthermore, there might also exist some time-varying shocks happening beyond the

[28]The data are obtained from Fan and Ying (2005).

[29]The total trade value and the number of college graduates are obtained from the *China Statistical Yearbook* (2003), and the number of SARS-infected persons is from Fan and Ying (2005).

province level that were correlated with the effect of the PHE in 2002 and affected the change in outcome variables. To address this concern, we divide provinces into different regions according to conventional criteria, that is, eastern, central, and western China.[30] Then, we include in the regression interactions between the region dummies and the marriage-year dummies. The results are shown in Panel B of Table 5, and we can see that they remain robust.

Permutation Test. Given that our sample comprises only 25 provinces, to check whether the variation in the PHE rate across provinces is sufficient to allow causal inference, a permutation test is conducted. To do so, we randomly assign prereform PHE ratios to provinces and estimate the effect of the "fake" treatment status on SWB. We repeat this exercise 500 times and plot the distributions of these coefficients in Figure 4. Fewer than 5% of the estimates are more negative than our estimates in Table 4, confirming that our main results are not driven by random factors.

Other Robustness Checks. In addition to the above three robustness checks, we also conduct several other checks to reinforce our results. (1) We investigate whether our results are robust to different specifications. First, we include the set of individual-level predetermined variables in Table 3 in the baseline regressions. The results remain robust. Second, we replace county fixed effects with province fixed effects, and the results remain robust. Third, we interact the province-level predetermined variables with a set of year-of-marriage dummies instead of a post 2003 dummy to enable a more flexible time-varying effect on SWB. The results are robust as well. (2) We include individuals married in 2003 in the sample, which are dropped in the main analysis to avoid the mixture of policy effects. The results remain similar. (3) We investigate whether the repeal of the PHE also affects other welfare measures such as emotional well-being. We use three variables to measure emotional well-being: the frequencies of feeling depressed, feeling nervous, and feeling restless. We

[30]Eastern China includes Beijing, Tianjin, Hebei, Liaoning, Shanghai, Jiangsu, Zhejiang, Fujian, Shandong, Guangdong and Hainan provinces. Central China includes Shanxi, Anhui, Jiangxi, Henan, Hubei, Hunan, Heilongjiang, and Jilin provinces. Western China includes Inner Menggu, Guangxi, Chongqing, Sichuan, Guizhou, Yunnan, Tibet, Shaanxi, Gansu, Qinghai, Ningxia, and Xinjiang provinces.

find that the PHE repeal resulted in a significant increase in the frequency of depression, nervousness, and restlessness, reconfirming our main results. (4) We conduct the same exercise using the urban sample and find no effect of the repeal of the PHE. (5) We calculate the standard errors using two other methods: by clustering over provinces and using a wild bootstrapping procedure and by clustering over counties. The results remain significant.

Due to space limitations, we present the aforementioned robustness checks in the Online Appendix Part C.

7 Channels: Sorting and Child Health

Hypothesis 2 and Hypothesis 3 predict that the repeal of the compulsory PHE reduced sorting on health (the health marker), which resulted in a decline in the expected health of children and increased sorting on SES. In this section, we investigate these channels.

7.1 Sorting by Health

To assess the change in assortative matching on health, we estimate the effect of the repeal of the compulsory PHE on the health gap between spouses.[31] Unfortunately, the CFPS does not contain information on health status before marriage, except for a small proportion who reported their birth weight. Thus, we must rely mainly on the health measurement at the time of the survey, including whether the respondent felt sick in the previous two weeks, whether he or she had been hospitalized in the past year, and self-reported health status. By doing so, we implicitly assume that current health status is correlated with past health. This may lead to downward bias in the measurement of the spousal gap in health given that spousal health habits and status tend to evolve in a similar way (Venters et al., 1984; Falba and Sindelar, 2008; Meyler et al., 2007; Pai et al., 2010). The results are reported in Table 6. Albeit underestimated, the estimates show that the husband–wife gap in recent illness and self-reported health increased significantly after the abolition of compulsory PHE,

[31] Similarly, Han et al. (2015) use the education gap and age gap between spouses to measure assortative matching on education and age, respectively.

supporting the model prediction that the repeal of compulsory PHE reduced PAM on health. [32]

7.2 Child Health

As spousal health is complementary in producing a healthy child, a decrease in sorting on health likely reduces overall child health.

We investigate how the PHE repeal affected the health of children who possess valid information pertaining to their health status,[33] by applying the DID model to estimate the effect on birth weight, height-for-age Z score, weight-for-age Z score, frequency of being sick in the past month, and frequency of seeing a doctor in the past month.[34] The regression results are reported in Table 7. The results consistently show that the PHE repeal tended to worsen child health. In provinces where the PHE rates were previously high, the decrease in children's birth weight and height and weight for their age and the increase in the likelihood of falling sick were larger than those in provinces where the PHE rates were low. The effects are statistically significant.

7.3 Matching on SES

As Hypothesis 3 predicts, the sorting tradeoff is likely to shift toward sorting on SES as the health cue becomes noisier. To test this prediction, we first use the respondent's own education as a proxy for SES and then turn to the fathers' and mothers' own education, following Sun and Zhang (2020), who argue that fathers' education is a better proxy for one's family SES than is one's own education in contemporary China. Interestingly, the

[32]Note that the survey did not cover each individual within a couple, which means that some observations corresponding to respondents whose partner was not in the survey are not included in the regression.

[33]We can match approximately 86.4% adults in the baseline sample with their child's information in the survey.

[34]We resort to the 2010 and 2012 waves to fill in the missing values on birth weight in wave 2014. We obtain the standard for calculating the height-for-age Z score under age 19 and weight-for-age Z score under age 10 from the WHO. Height- and weight-for-age Z scores are extensively used in the literature as measures of child health (see Thomas et al., 1991, Strauss and Thomas, 1998, and Chen and Li, 2009, for example).

spousal gap in both one's own and father's education years decreased more in provinces with higher ex ante PHE rates after the repeal (Columns (1) and (2) of Table 8), suggesting that the repeal increased sorting on SES. Our findings are consistent with our model predictions.

8 Testing Inequality Implications

Our model predicts that the repeal of the PHE is more likely to hurt the healthy, the poor and women (Hypothesis 4). We test these implications for inequality in this section.

8.1 By Health Status

We first examine the heterogeneity in the treatment effect by health status. Given that spousal health is complementary in marital production, healthy individuals would suffer more as the repeal of the PHE reduced marital sorting by health.

Since we have no information on premarital health, we use recent illness at the time of the survey as a proxy.[35] In doing so, we assume that the respondent's health remains stable for a certain period. In Table 9, we find statistically significant effects for the healthy subsample, while most of the estimates for the unhealthy subsample are statistically insignificant and smaller in magnitude. The results show that healthy people suffered more loss of SWB from the PHE repeal, consistent with our model prediction. This result is similar in spirit to the results in Angelucci and Bennett (2021), who find that frequent HIV testing significantly increases the probabilities of marriage and pregnancy for people not at risk while decreasing them for people who are more at risk.

8.2 By SES

We examine the heterogeneity in the policy effect by SES in this section. We divide the sample into low-SES and high-SES groups based on the respondent's (or his or her father's)

[35]Fewer than 30% of the individuals in our sample reported birth weights, so birth weight is not a suitable metric for dividing subsamples. People who felt sick in the previous two weeks are defined as being unhealthy, and those whose did not are defined as being healthy.

educational attainment—individuals (or their fathers) with a junior high school diploma and below are classified into the low-SES group and the others belong to the high-SES group. We examine the policy effect on life satisfaction, family satisfaction, and happiness for the two groups. The results are reported in Table 10. Panel A uses the respondent's own education to divide the subsample, and Panel B uses the father's education to divide the subsample. Across the three measures of SWB, the deterioration effects are consistently salient only for the low-SES group, supporting our model prediction. The empirical findings suggest that the benefit to less healthy people from the PHE repeal does not offset the negative effect on other people in the low-SES group.

8.3 By Gender

We investigate whether the effects differ between men and women. In reality, mothers are typically the main caregivers, especially for children (e.g., Han and Shi, 2019). If we extend our model to allow the within-couple transfer not to be fully flexible, a deterioration in child health tends to have a greater negative impact on females. Therefore, it predicts that women, on average, will be worse off as the health cue becomes noisier. We estimate the impact on SWB for male and female respondents separately. The results in Table 11 confirm that the negative effect of the PHE repeal on SWB is more salient for women.

9 Conclusion

Becker (1981) notes that information friction on partner traits gives rise to search costs in the marriage market and influences sorting patterns. This paper builds upon this notion and takes a step further by examining the impact of information noise on marital matching and postmarital well-being both theoretically and empirically. We take advantage of the repeal of the compulsory PHE in China, which increased noise in the health cue, as a natural experiment to examine how noise in matching affects SWB in marital life by distorting matching patterns. Using a DID strategy, we find that the removal of the compulsory PHE resulted in a significant drop in SWB after marriage by reducing positive assortative mating on health and increasing positive assortative mating on SES. The change in the matching pattern leads to poorer child health, which is likely a major channel for the deterioration of SWB. In particular, the healthy, the poor and women tended to suffer more in their marriage payoffs from the repeal of the PHE.

Despite the associated ethical controversy, the compulsory PHE has been strongly advocated by professionals from the WHO (Rennie and Mupenda, 2008). The findings of this paper provide supporting evidence for the positive impact of the PHE on marital utility and, especially, child health. In addition, our analysis suggests that given the complementarity of spousal traits in producing marital surplus, making health information less transparent would reduce overall welfare. A possible policy direction to protect the underprivileged is to reduce complementarity by reducing the impact of parental health on child health or reducing the health burden on families rather than suppressing information.

Our findings also have important implications for inequality policies. A tendency in attempts to address inequality in college admissions and marriage and labor markets is to suppress information to avoid competition based on explicit measures. One recent example is that the University of California system extended its policy of test-free admissions in 2021 to resolve a 2019 lawsuit charging that the SAT and ACT are biased against poor, Black and Hispanic students. Such measures are not rare in health care and labor markets. However, our analysis shows that this type of attempt may risk exacerbating inequality along different dimensions even when matching efficiency is compromised. In particular, concealing information cues on some dimensions tends to shift the sorting tradeoff toward wealth, which is often at odds with the policy intention.

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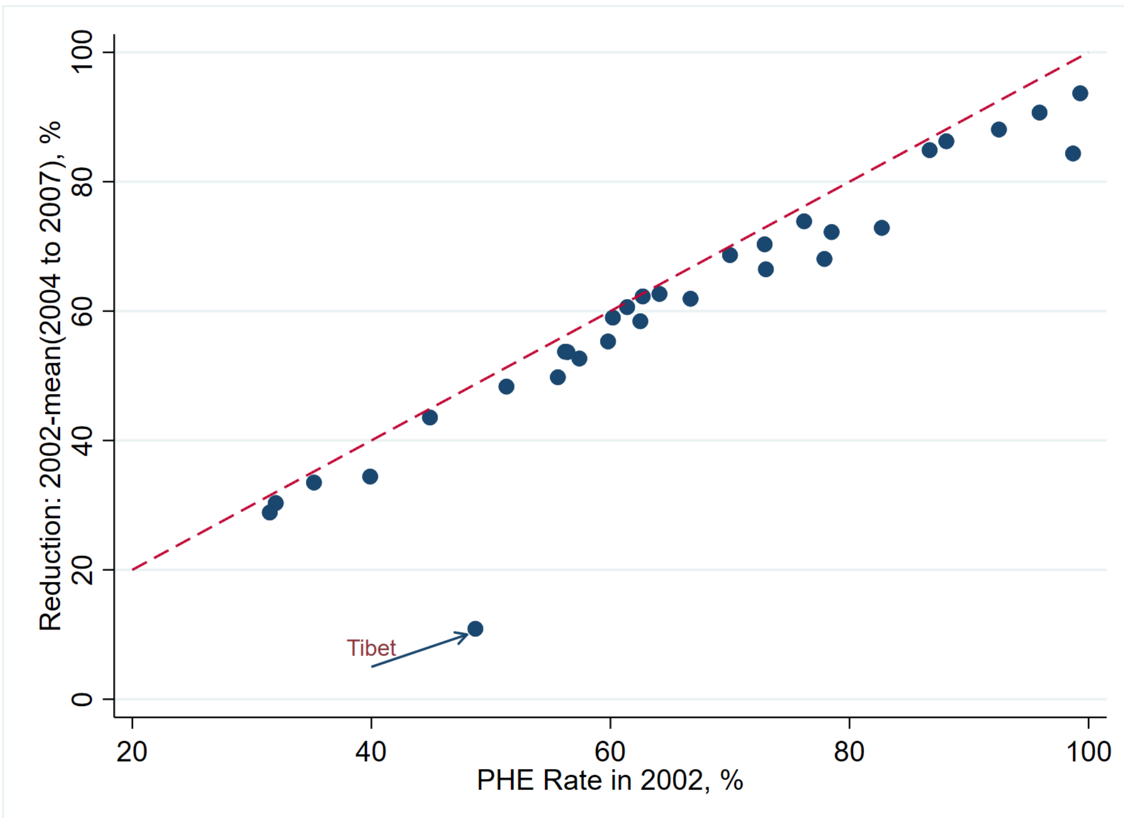
Figure 1. PHE Rate between 1998 and 2009

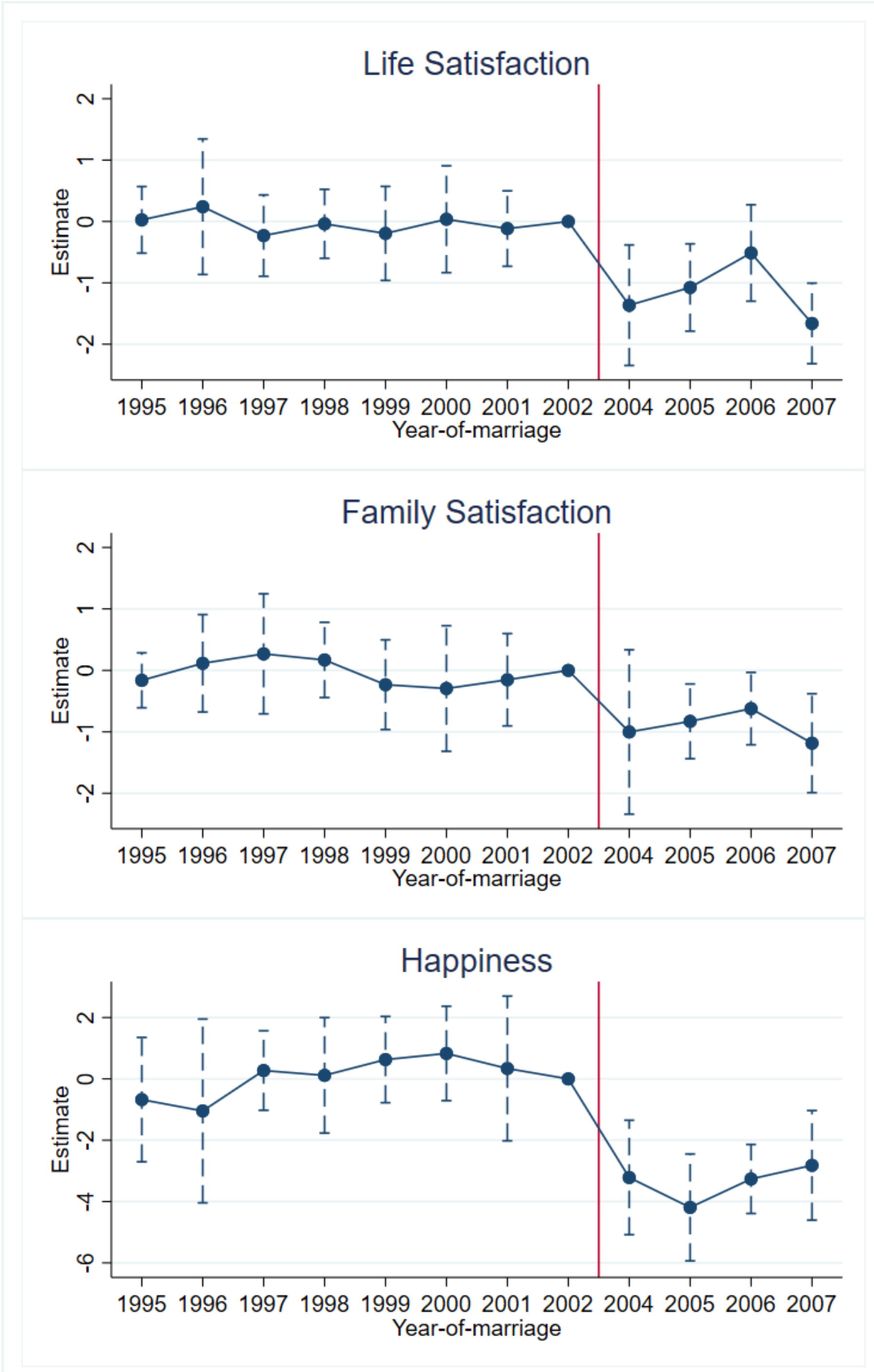
Notes: The dots represent the national average PHE rate in China from 1998 to 2009. The data are from the *China Health Statistical Yearbook*.

Figure 2. Relationship between Reductions in the PHE Rate after 2003 and the Initial Level in 2002



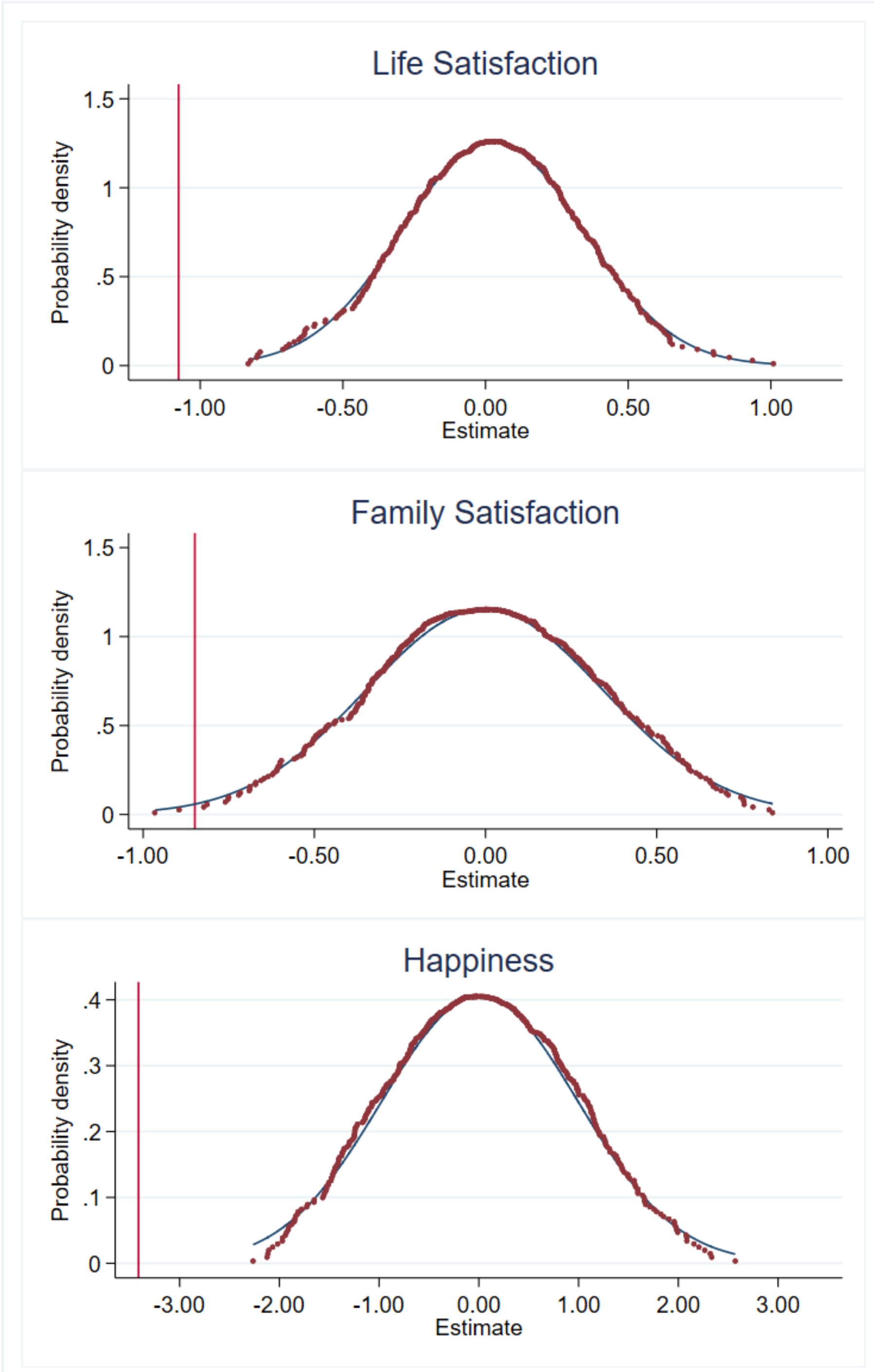
Notes: The dots represent the observations for each province. The dashed line is the 45-degree line. The data are obtained from the *China Health Statistical Yearbook*.

Figure 3. Parallel Trend



Notes: The omitted benchmark for year of marriage is 2002. The sample does not include individuals who married in 2003. The points represent estimated coefficients β_t from Equation (2), and vertical dashed bars represent 95% confidence intervals.

Figure 4. Distribution of Estimates of the Permutation Test



Notes: The figures show the probability distribution density of the coefficients of the permutation test, which randomly assigns the treatment intensity to the provinces 500 times. The vertical lines represent the baseline coefficients in Table 4.

Table 1. Summary Statistics

Variable	Mean	SD	N
Panel A: Individual-level variables			
Life satisfaction	3.768	1.015	3,213
Family satisfaction	3.864	0.994	3,213
Happiness	7.449	2.216	3,213
Ethnic minority	0.118	0.323	3,212
Gender	0.482	0.500	3,213
Age	37.628	6.612	3,213
Number of siblings	2.523	1.569	2,614
Education level	2.361	0.973	3,213
Father's education level	2.001	0.974	2,978
Mother's education level	1.464	0.743	3,037
Log(weeks not living with father before age 12+1)	0.441	1.206	2,525
Log(weeks not living with mother before age 12+1)	0.225	0.924	2,575
Panel B: Province-level variables in 2002			
PHE rate	0.689	0.180	25
Log(GDP per capita)	9.115	0.617	25
Log(consumption per capita)	8.209	0.499	25
Share of primary industry in GDP	0.148	0.065	25
Health institutes per 10,000 people	2.494	0.889	25
Doctors per 10,000 people	15.499	7.551	25
Civil affairs staff per 10,000 people	0.913	0.241	25
Log(government health expenditures per capita)	3.882	0.596	25
Infectious diseases infection rate	18.982	6.520	25

Table 2. Impact of the Policy on Provincial Marriage Rates

	(1)
Variable	Marriage rate (%)
PHE rate×Post2003	-0.029
	(0.208)
Observations	347
R-squared	0.732
Province control	Yes
Province FE	Yes
Year-of-Marriage FE	Yes

Notes: The regression function is $Marriage\_rate_{pt} = \beta \times PheRate_p \times Post2003_t + \psi_p \times Post2003_t + \delta_p + \delta_t + \epsilon_{pt}$, in which t denotes the calendar year rather than the year of marriage and δ_p denotes province fixed effects. Data on the marriage rate are obtained from the China Civic Affairs Statistical Yearbook. The standard errors are reported in parentheses, two-way clustered by province and year of marriage.

Table 3. Balance Test of Predetermined Individual Characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Variables	Father's education	Mother's education	Education	Number of siblings	Ethnic minority	Log(weeks not living with father before age 12+1)	Log(weeks not living with mother before age 12+1)
Mean of y	2.001	1.464	2.361	2.523	0.118	0.441	0.225
PHE rate×Post2003	0.498	0.120	0.776	-0.627	0.005	0.292	0.140
	(0.810)	(0.427)	(0.574)	(0.797)	(0.123)	(1.109)	(0.782)
Observations	2,978	3,037	3,213	2,614	3,212	2,525	2,575
R-squared	0.193	0.209	0.287	0.268	0.641	0.126	0.109
Province control	Yes	Yes	Yes	Yes	Yes	Yes	Yes
County FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year-of-Marriage FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The standard errors are reported in parentheses, two-way clustered by province and year of marriage.

Table 4. Impact of the Policy on Subjective Well-Being

	(1)	(2)	(3)	(4)	(5)	(6)
VARIABLES	Life Satisfaction		Family Satisfaction		Happiness	
Mean of y	3.768		3.864		7.449	
PHE rate×Post2003	-0.916***	-1.076***	-0.976***	-0.849***	-1.716***	-3.412***
	(0.114)	(0.269)	(0.038)	(0.271)	(0.183)	(0.489)
Observations	3,213	3,213	3,213	3,213	3,213	3,213
R-squared	0.120	0.122	0.124	0.125	0.130	0.132
Province control	No	Yes	No	Yes	No	Yes
County FE	Yes	Yes	Yes	Yes	Yes	Yes
Year-of-Marriage FE	Yes	Yes	Yes	Yes	Yes	Yes

Notes: The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 5. Robustness Check on Concurrent Events

	(1) Life Satisfaction	(2) Family Satisfaction	(3) Happiness
Panel A: Inclusion of confounding shocks			
PHE rate×Post2003	-1.132***	-1.027**	-3.575***
	(0.243)	(0.375)	(0.492)
Observations	3,213	3,213	3,213
R-squared	0.123	0.126	0.132
Confounding Shocks	Yes	Yes	Yes
Panel B: Inclusion of region×year-of-marriage fixed effect			
PHE rate×Post2003	-1.089***	-1.001***	-3.639**
	(0.300)	(0.131)	(1.235)
Observations	3,213	3,213	3,213
R-squared	0.128	0.132	0.139
Region×year-of-marriage FE	Yes	Yes	Yes

Notes: We include provincial controls and county and year-of-marriage fixed effects in all regressions. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively. Confounding shocks include trade value, number of college graduates and SARS-infected persons in each province in 2002, interacted with a post-2003 dummy.

Table 6. Channel: Assortative Matching of Health

	(1)	(2)	(3)	(4)
	Gap in birth weight	Gap in recent illness	Gap in hospitalization	Gap in self-reported health status
Mean of y	0.868	0.318	0.129	1.033
PHE rate×Post2003	1.388	0.712**	0.010	1.967**
	(4.019)	(0.275)	(0.397)	(0.855)
Observations	188	1,203	1,073	1,203
R-squared	0.388	0.155	0.147	0.149

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 7. Channel: Child Health

	(1)	(2)	(3)	(4)	(5)
	Birth weight	Height-for-age Z score	Weight-for-age Z score	Frequency of sickness in the last month	Frequency of seeing a doctor in the last month
Mean of y	6.401	-1.660	-0.267	0.414	0.325
PHE rate×Post2003	-0.947**	-3.278*	-1.746**	0.685**	0.156
	(0.323)	(1.700)	(0.760)	(0.259)	(0.154)
Observations	2,481	2,793	1,847	2,939	2,942
R-squared	0.148	0.233	0.176	0.141	0.112

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 8. Channel: Assortative Matching on SES

	(1) Gap in education years	(2) Gap in father's education years	(3) Gap in mother's education years
Mean of y	2.587	3.359	2.319
PHE rate×Post2003	-4.543**	-4.401*	-1.199
	(2.027)	(2.240)	(2.180)
Observations	1,031	1,046	1,080
R-squared	0.165	0.189	0.211

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 9. Heterogeneity Test by Health Status

	(1)	(2)	(3)	(4)	(5)	(6)
	Life Satisfaction		Family Satisfaction		Happiness	
	Healthy	Unhealthy	Healthy	Unhealthy	Healthy	Unhealthy
PHE rate×Post2003	-1.281**	-0.940***	-1.222***	-0.008	-3.897***	-2.945
	(0.430)	(0.239)	(0.367)	(0.325)	(0.698)	(2.703)
Observations	2,445	767	2,445	767	2,445	767
R-squared	0.145	0.237	0.150	0.221	0.149	0.276

Notes: People who felt sick in the previous two weeks are defined as unhealthy, and those whose did not are defined as healthy. We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 10. Heterogeneity Test by SES

	(1)	(2)	(3)	(4)	(5)	(6)
	Life Satisfaction		Family Satisfaction		Happiness	
Panel A: Own Education	Low	High	Low	High	Low	High
PHE rate×Post2003	-1.615***	-0.536	-1.095**	-0.533	-5.317***	-0.702
	(0.501)	(0.723)	(0.434)	(0.568)	(0.889)	(1.859)
Observations	1,712	1,501	1,712	1,501	1,712	1,501
R-squared	0.146	0.209	0.154	0.206	0.153	0.209
Panel B: Father's Education	Low	High	Low	High	Low	High
PHE rate×Post2003	-1.727***	0.715	-0.901***	-0.026	-5.386***	1.820
	(0.292)	(0.827)	(0.243)	(1.001)	(1.081)	(1.504)
Observations	2,163	1,050	2,163	1,050	2,163	1,050
R-squared	0.145	0.230	0.141	0.231	0.149	0.236

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table 11. Heterogeneity Test by Gender

	(1)	(2)	(3)	(4)	(5)	(6)
	Life Satisfaction		Family Satisfaction		Happiness	
	Female	Male	Female	Male	Female	Male
PHE rate×Post2003	-1.662***	-0.666	-1.389***	-0.541	-4.842***	-2.144
	(0.270)	(0.857)	(0.231)	(0.590)	(1.108)	(1.288)
Observations	1,664	1,549	1,664	1,549	1,664	1,549
R-squared	0.168	0.177	0.161	0.182	0.185	0.175

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Appendices to:
Ignorance is Whose Bliss: The Repeal of Compulsory Premarital Health Examinations and Marital Outcomes in Rural China

Li Han, Xinzheng Shi, and Ming-ang Zhang

First Version, May 2021; This Version, November 2023

A Model

A.1 Setup

Our transferable utility (TU) model describes a marriage market where individuals of both genders possess two fixed binary traits: health and socioeconomic status (SES). We describe men and women using the notation (x, X) and (y, Y) respectively. x and y indicate the health status for Mr. (x, X) and Ms. (y, Y), $x, y \in \{0, 1\}$ with 0 representing low health status and 1 representing high health status. SES status is denoted as X and Y, $X, Y \in \{p, r\}$ with values "p" for low SES status and "r" for high SES status, respectively, for Mr. (x, X) and Ms. (y, Y).

Therefore there are four types of individuals in each of the two genders, namely types $1r$, $1p$, $0r$, $0p$. The measures for each type of the two genders are assumed as follows

	$1r$	$1p$	$0r$	$0p$
M	1	δ_1	1	δ_2
F	μ	$\delta_1\mu$	μ	$\delta_2\mu$

where $0 < \delta_1 \leq \delta_2$ and assumption 1 holds.

Assume that the total utility from the marriage between Mr. (x, X) and Ms. (y, Y) is

$$b(x, X) + g(y, Y) = U_{xX,yY} = q(x, y)t(X, Y)$$

where $q(1, 1) = 1,\ q(0, 1) = q(1, 0) = \lambda,\ q(0, 0) = \lambda^2,\ 0 < \lambda < 1$.

For convenience, denote $\tilde{\delta_1} = \frac{\delta_1}{\delta_1+\delta_2}$, $\tilde{\delta_2} = \frac{\delta_2}{\delta_1+\delta_2}$, and $\tilde{\lambda} = \tilde{\delta_1} + \tilde{\delta_2}\lambda$.

We focus on the case where $\Delta t^{rp}_{pp} - \lambda\Delta t^{rr}_{rp} > 0$ and $\tilde{\lambda}/\lambda < t^{rp}/t^{pp}$. Here, Δt^{rp}_{pp} is defined as $t(r, p) - t(p, p)$ and Δt^{rr}_{rp} is defined similarly.

A.2 Proofs

We begin by the following lemma.

Lemma 1. *In each health–SES–gender group, all individuals have the same expected utility in equilibrium.*

Proof. Suppose that i and i' both belong to the same health-SES-gender group, with $u_i^* > u_{i'}^*$ in an equilibrium. Because of the first condition in Definition 1, i must be matched to a spouse j. Suppose j obtains u_j^* in the equilibrium. We must have $u^{i,j} + v^{i,j} = u_i^* + v_j^*$. Because $u^{i,j} + v^{i,j} = u^{i',j} + v^{i',j}$, we have $u^{i',j} + v^{i',j} > u_{i'}^* + v_j^*$, which contradicts the second condition in Definition 1. □

This lemma is intuitive. If there are two identical individuals obtaining different equilibrium expected utilities, the one receiving the higher expected utility must be married. The one who receives the lower expected utility can always outbid the individual who receives higher expected utility by slightly increasing the transfer to the high-utility person's spouse without making himself (or herself) worse off. Therefore, we can simplify our analysis by focusing on the types of individuals instead of on the individuals themselves.

A.2.1 Proof of Propositions

Case 1: Both traits are observable.

We can first show the following lemma.

Lemma 2. *In equilibrium, the set of single persons must include some low-poor men but must never include a woman or a high-rich man.*

Proof. Because there are more men than women, for any equilibrium there always exists a man xX who remains single. Suppose there is also a woman yY remaining single in the equilibrium. Because of complementarity (Assumption 2), $t(X,Y) = t(X,Y) + t(\phi_X, \phi_Y) \geq t(X, \phi_Y) + t(\phi_X, Y)$. So i and j as a deviating couple can get at least $t(X,Y)$, violating condition 2 of Definition 1. So all women should be married in any equilibrium.

Suppose there is a type 1r man who remains single in an equilibrium and receives $t(r, \phi_Y)$. Because $\frac{2+\delta_1}{2+\delta_1+\delta_2} < \mu < 1$ and all women are married in equilibrium, at least one man $xX \in \{1p, 0r, 0p\}$ is married to a woman $yY \in \{1p, 0r, 0p\}$. Suppose in equilibrium xX (yY) gets Eu_{xX}^* (Ev_{yY}^*) respectively. Because of Definition 1, we must have $Ev_{yY}^* = Eq(x,y)t(X,Y) - Eu_{xY}^* \leq Eq(x,y)t(X,Y) - t(X, \phi_Y)$. Because of Assumption 2, $t(X,Y) -$

$t(X, \phi_Y) \leq t(X, Y) - t(X, \phi_Y)$ and $q(0, y) < q(1, y)$. So $Ev^*_{yY} < Eq(1, y)t(r, Y) - t(r, \phi_Y)$. Re-arranging terms, we have $Ev^*_{yY} + t(r, \phi_Y) < Eq(1, y)t(r, Y)$, which contradicts condition 2 of Definition 1. So all high-rich men should be married in any equilibrium. Similarly we can show that men of types $1p$ and $0r$ are also married in the equilibrium Given the above results and Assumption 2, some low-poor men (type $0p$) have to remain single. $\square$

Given Lemmas 1 & 2, we can show that the equilibrium marriage assignment is summarized in the following matrix:

$$\mathbf{A} = \left[\begin{array}{c|ccccc} & 1r & 1p & 0r & 0p & \phi \\ \hline 1r & \mu & 1-\mu & & & \\ 1p & & \delta_1\mu + \mu - 1 & (1+\delta_1)(1-\mu) & & \\ 0r & & & \mu - (1+\delta_1)(1-\mu) & (2+\delta_1)(1-\mu) & \\ 0p & & & & \delta_2\mu - (2+\delta_1)(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \\ \phi & & & & & \end{array}\right]$$

where the element in the matrix $A_{i,j}$ represents the size of type-i men who are matched with type-j women, $i, j \in \{1r, 1p, 0r, 0p\}$.

The equilibrium expected payoffs of men (u^*_i) and of women (v^*_j) are summarized below.

$$Eu^*_i = \begin{cases} \Delta t^{rp}_{pp} + \lambda t^{pr} - \lambda^2(\Delta t^{rr}_{rp} + t^{pp}) & \text{if } i = 1r \\ \lambda t^{rp} - \lambda^2(\Delta t^{rr}_{rp} + t^{pp}) & \text{if } i = 1p \\ \lambda^2 \Delta t^{rp}_{pp} & \text{if } i = 0r \\ 0 & \text{if } i = 0p \end{cases}$$

$$Ev^*_j = \begin{cases} (1+\lambda^2)(\Delta t^{rr}_{rp} + t^{pp}) - \lambda t^{pr} & \text{if } j = 1r \\ t^{pp} - \lambda t^{pr} + \lambda^2(\Delta t^{rr}_{rp} + t^{pp}) & \text{if } j = 1p \\ \lambda^2(\Delta t^{rr}_{rp} + t^{pp}) & \text{if } j = 0r \\ \lambda^2 t^{pp} & \text{if } j = 0p \end{cases}$$

It is straightforward that the above assignments and payoffs satisfy conditions 1 and 2 of Definition 1 (the definition of the market equilibrium). Next we show that condition

3 of the equilibrium is also satisfied, $Eu_i^* + Ev_j^* \geq Eu^{i,j} + Ev^{i,j}$, for any i, j. That is, the equilibrium cannot be blocked by a deviating individual (couple).

1. $i = 1r,\ j = 0r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \; & \Delta t_{pp}^{rp} + \lambda t^{pr} - \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) + \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{rr} \\ = \; & \Delta t_{pp}^{rp} - \lambda \Delta t_{rp}^{rr} > 0 \end{aligned}$$

2. $i = 1p,\ j = 1r$

$$\begin{aligned} Eu_i^* + Ev_j^* - EU_{i,j} &= \lambda t^{rp} - \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) + (\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{pr} + \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) - t^{rp} \\ &= \Delta t_{rp}^{rr} - \Delta t_{pp}^{pr} > 0 \end{aligned}$$

3. $i = 1p,\ j = 0p$

$$\begin{aligned} Eu_i^* + Ev_j^* - EU_{i,j} &= \lambda t^{rp} - \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) + \lambda^2 t^{pp} - \lambda t^{pp} \\ &= \lambda(\Delta t_{pp}^{rp} - \lambda \Delta t_{pr}^{rr}) \geq 0 \end{aligned}$$

4. $i = 0r,\ j = 1r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \; & \lambda^2 \Delta t_{pp}^{rp} + (\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{pr} + \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{rr} \\ = \; & (1+\lambda^2)t^{rr} - \Delta t_{pp}^{rp} - \lambda t^{rp} - \lambda t^{rr} = t^{rr} - \Delta t_{pp}^{rp} + \lambda^2 t^{rr} - \lambda t^{rp} - \lambda t^{rr} \\ > \; & t^{rp} + \lambda^2 t^{rr} - \lambda t^{rp} - \lambda t^{rr} = (1-\lambda)[t^{rp} - \lambda t^{rr}] > 0 \end{aligned}$$

because $\lambda < \Delta t_{pp}^{rp}/\Delta t_{rp}^{rr} < t^{rp}/t^{rr} < t^{pp}/t^{rp}$ (the second and third inequalities due to log complementarity of t).

5. $i = 0r,\ j = 1p$

$$\begin{aligned} Eu_i^* + Ev_j^* - EU_{i,j} &= \lambda^2 \Delta t_{pp}^{rp} + t^{pp} - \lambda t^{pr} + \lambda^2(\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{pr} \\ &= t^{pp} + \lambda^2 t^{rr} - 2\lambda t^{rp} = t^{rr}(\lambda - \frac{t^{rp}}{t^{rr}})^2 + \frac{t^{rr}t^{pp} - (t^{rp})^2}{t^{rr}} \geq 0 \end{aligned}$$

because of the log complementarity in t.

6. $i = 0p,\ j = 1r$

$$\begin{aligned} Eu_i^* + Ev_j^* - EU_{i,j} &= (1+\lambda^2)(\Delta t_{rp}^{rr} + t^{pp}) - \lambda t^{rp} - \lambda t^{rp} \\ &\geq (1+\lambda^2)t^{rp} - 2 - \lambda t^{rp} = (1-\lambda)^2 t^{rp} > 0 \end{aligned}$$

7. For $i = 0p$ or, $j = 0p$, the way that the equilibrium payoff is derived guarantees that

$$Eu_i^* + Ev_j^* - EU_{i,j} \geq 0$$

Note that the measure of mixed marriages across the wealth status is

$$A(mixed\ wealth) = A_{1r,1p} + A_{1p,0r} + A_{0r,0p} = 2(2+\delta_1)(1-\mu)$$

The measure of mixed marriages across the health marker is

$$A(mixed\ marker) = A_{1p,0r} = (1-\mu)(1+\delta_1)$$

The expected health of children

$$Eq^* = \mu(1+\delta_1) + (1-\mu)(1+\delta_1)\lambda + [(1+\delta_2)\mu - (1+\delta_1)(1-\mu)]\lambda^2$$

<u>Case 2</u>: unobservable health status but type $1r$ can signal

This case occurs when the signalling cost for type $1r$ is small enough while that for type $1p$ is big enough. Thus types 1p and 0p would be indistinguishable ex ante. Lemma 2 can can be adjusted in this case to the following: In equilibrium, the set of single persons must include some poor men but must never include a woman or a high-rich man. The equilibrium marriage assignment is summarized in the following matrix:

$$\mathbf{A} = \left[\begin{array}{c|cccc} & 1r & 0r & p & \phi \\ \hline 1r & \mu & 1-\mu & & \\ 0r & & 2\mu - 1 & 2(1-\mu) & \\ p & & & (\delta_1+\delta_2)\mu - 2(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \\ \phi & & & & \end{array}\right]$$

where $A_{i,j}$ equals to the size of type-i men who are matched with type-j women.

The equilibrium expected payoffs of men (u_i^*) and of women (v_j^*) are summarized below.

$$Eu_i^* = \begin{cases} \lambda(1-\lambda)t^{rr} + \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) & \text{if } i = 1r \\ 0 & \text{if } i = 1p \\ \lambda\tilde{\lambda} t^{rp} - \tilde{\lambda}^2 t^{pp} & \text{if } i = 0r \\ 0 & \text{if } i = 0p \end{cases}$$

$$Ev_j^* = \begin{cases} (1-\lambda+\lambda^2)t^{rr} - \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) & \text{if } j = 1r \\ \tilde{\lambda} t^{pp} & \text{if } j = 1p \\ \lambda^2 t^{rr} - \lambda\tilde{\lambda} t^{rp} + \tilde{\lambda}^2 t^{pp} & \text{if } j = 0r \\ \lambda\tilde{\lambda} t^{pp} & \text{if } j = 0p \end{cases}$$

where $\tilde{\lambda} = \tilde{\delta}_1 + \tilde{\delta}_2\lambda$.

It is straightforward that the above assignments and payoffs satisfy conditions 1 and 2 of Definition 1 (the definition of the market equilibrium). Next we show that condition 3 of the equilibrium is also satisfied, $Eu_i^* + Ev_j^* \geq Eu^{i,j} + Ev^{i,j}$, for any i, j. That is, the equilibrium cannot be blocked by a deviating individual (couple).

1. $i = 1r,\ j = p$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \;& \lambda(1-\lambda)t^{rr} + \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) + \tilde{\lambda}^2 t^{pp} - \tilde{\lambda} t^{rp} \\ = \;& \lambda(1-\lambda)t^{rr} + \tilde{\lambda}\lambda t^{pr} - \tilde{\lambda} t^{rp} = (1-\lambda)[\lambda t^{rr} - \tilde{\lambda} t^{rp}] > 0 \end{aligned}$$

   if $\tilde{\lambda}/\lambda < t^{rr}/t^{rp}$. It holds under Assumption 3.

2. $i = 0r,\ j = 1r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \;& \lambda\tilde{\lambda} t^{rp} - \tilde{\lambda}^2 t^{pp} + (1-\lambda+\lambda^2)t^{rr} - \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) - \lambda t^{rr} = (1-\lambda)^2 t^{rr} > 0 \end{aligned}$$

3. $i = p,\ j = 1r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \;& (1-\lambda+\lambda^2)t^{rr} - \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) - \tilde{\lambda} t^{pr} \end{aligned}$$

Let $G = Eu_i^* + Ev_j^* - EU_{i,j}$ in this case. It is easy to show that $\frac{\partial G}{\partial \tilde{\lambda}} < 0$. We know $\tilde{\lambda} \leq \frac{1+\lambda}{2}$ as $\tilde{\delta_1} \leq 1/2$. We can show that

$$G|_{\tilde{\lambda}=\frac{1+\lambda}{2}} = \frac{3}{4}(1-\lambda)^2 t^{rr} + \frac{(1+\lambda)^2}{4}(t^{rr} - 2t^{rp} + t^{pp}) > 0$$

So $G > 0$ for $\tilde{\lambda} \leq \frac{1+\lambda}{2}$.

4. $i = p,\ j = 0r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \ & \lambda^2 t^{rr} - \lambda\tilde{\lambda}t^{rp} + \tilde{\lambda}^2 t^{pp} - \lambda\tilde{\lambda}t^{pr} \\ = \ & \lambda(\lambda t^{rr} - \tilde{\lambda}t^{rp}) - \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda}t^{pp}) \geq 0 \end{aligned}$$

under the assumption of log complementarity of t.

Note that the measure of mixed marriages across the wealth status is

$$A(mixed\ wealth) = A_{0r,p} = 2(1-\mu)$$

The measure of mixed marriages across the health status is

$$\begin{aligned} A(mixed\ health) &= A_{1r,0r} + \tilde{\delta_1}A_{0r,p} + 2\tilde{\delta_1}\tilde{\delta_2}A_{p,p} \\ &= (1-\mu)(1+2\tilde{\delta_1} - 4\tilde{\delta_1}\tilde{\delta_2}) + 2\tilde{\delta_1}\delta_2\mu \\ &= (1-\mu)(1-2\tilde{\delta_1}) + 2\tilde{\delta_1}\delta_2\mu \end{aligned}$$

The expected health of children

$$Eq^* = \mu + (1-\mu)\lambda + (2\mu - 1)\lambda^2 + 2(1-\mu)(\lambda\tilde{\lambda} - \tilde{\lambda}^2) + [(\delta_1 + \delta_2)\mu]\tilde{\lambda}^2$$

Case 3: unobservable health status but no one signals

This case occurs when the signalling cost is too high for both type $1r$ and type $1p$. The matching is effectively a one-dimensional matching problem where only the SES status is observable. Following the same spirit of Lemma 2, we can show that in equilibrium, the set of single persons must include some poor men but must never include a woman or a high-rich man. The matching along the health dimension is random. The equilibrium marriage assignment is summarized in the following matrix:

$$\mathbf{A} = \left[\begin{array}{c|ccc} & r & p & \phi \\ \hline r & 2\mu & 2(1-\mu) & \\ p & & (\delta_1+\delta_2)\mu - 2(1-\mu) & (2+\delta_1+\delta_2)(1-\mu) \\ \phi & & & \end{array}\right]$$

where $a_{i,j}$ equals to the size of type-i men who are matched with type-j women.

The equilibrium expected payoffs of men (u_i^*) and of women (v_j^*) are summarized below.

$$Eu_i^* = \begin{cases} \tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp} & \text{if } i = 1r \\ 0 & \text{if } i = 1p \\ \lambda\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp} & \text{if } i = 0r \\ 0 & \text{if } i = 0p \end{cases}$$

$$Ev_j^* = \begin{cases} \frac{1+\lambda}{2}t^{rr} - [\frac{1+\lambda}{2}\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp}] & \text{if } j = 1r \\ \tilde{\lambda}t^{pp} & \text{if } j = 1p \\ \frac{(1+\lambda)\lambda}{2}t^{rr} - [\frac{1+\lambda}{2}\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp}] & \text{if } j = 0r \\ \lambda\tilde{\lambda}t^{pp} & \text{if } j = 0p \end{cases}$$

where $\tilde{\lambda} = \tilde{\delta_1} + \tilde{\delta_2}\lambda$.

It is straightforward that the above assignments and payoffs satisfy conditions 1 and 2 of Definition 1 (the definition of the market equilibrium). Next we show that condition 3 of the equilibrium is also satisfied, $Eu_i^* + Ev_j^* \geq Eu^{i,j} + Ev^{i,j}$, for any i, j. That is, the equilibrium cannot be blocked by a deviating individual (couple).

1. $i = p,\ j = r$

$$\begin{aligned} & Eu_i^* + Ev_j^* - EU_{i,j} \\ = \ & \frac{(1+\lambda)^2}{4}t^{rr} - (1+\lambda)\tilde{\lambda}t^{pr} + \tilde{\lambda}^2 t^{pp} > 0 \end{aligned}$$

because of log complementarity.

Note that the measure of mixed marriages across the wealth status is

$$A(mixed\ wealth) = A_{0r,p} = 2(1-\mu)$$

The measure of mixed marriages across the health status is

$$A(mixed\ health) \quad = \quad 1 + [(\delta_1 + \delta_2)\mu - 2(1-\mu)]2\tilde{\delta}_1\tilde{\delta}_2$$

The expected health of children

$$Eq^* = \frac{1}{2}\mu(1+\lambda)^2 + (1-\mu)(1+\lambda)\tilde{\lambda} + \tilde{\lambda}^2[(\delta_1+\delta_2)\mu - 2(1-\mu)]$$

Proof of Proposition 1:

The measures for the mixed-wealth marriages and the mixed-health marriages can be respectively illustrated as below.

$$A(mixed\ SES) = \begin{cases} 2(2+\delta_1)(1-\mu) & \text{case 1} \\ 2(1-\mu) & \text{case 2} \\ 2(1-\mu) & \text{case 3} \end{cases}$$

$$A(mixed\ health) = \begin{cases} (1+\delta_1)(1-\mu) & \text{case 1} \\ (1-\mu)(1+2\tilde{\delta_1}) + [(\delta_1+\delta_2)\mu - 2(1-\mu)]2\tilde{\delta_1}\tilde{\delta_2} & \text{case 2} \\ 1 + [(\delta_1+\delta_2)\mu - 2(1-\mu)]2\tilde{\delta_1}\tilde{\delta_2} & \text{case 3} \end{cases}$$

It is obvious that $A^{(1)}(mixed\ SES) > A^{(2)}(mixed\ SES) = A^{(3)}(mixed\ SES)$. That is, mixed-SES marriages decreases from case 1 to case 3.

For mixed-health marriages, it is obvious that $A^{(2)}(mixed\ health) < A^{(3)}(mixed\ health)$ because $\mu \geq 3/4$ and $\tilde{\delta_1} \leq 1/2$;

$$\begin{aligned} & A^{(2)}(mixed\ health) - A^{(1)}(mixed\ health) \\ = \ & (1-\mu)(1+2\tilde{\delta_1}) + [(\delta_1+\delta_2)\mu - 2(1-\mu)]2\tilde{\delta_1}\tilde{\delta_2} - (1+\delta_1)(1-\mu) \\ = \ & (1-\mu)(2\tilde{\delta_1} - \delta_1 - 4\tilde{\delta_1}\tilde{\delta_2}) + \delta_1\mu 2\tilde{\delta_2} \\ \geq \ & -(1-\mu)[\delta_1 - 2\tilde{\delta_1}(\tilde{\delta_2} - \tilde{\delta_1})] + \mu\delta_1 \\ > \ & \mu\delta_1 - (1-\mu)\delta_1 > 0 \end{aligned}$$

Therefore, $A^{(1)}(mixed\ health) < A^{(2)}(mixed\ health) < A^{(3)}(mixed\ health)$. In other words, the sorting on wealth increases and the sorting on health decreases. E.O.P.

Proof of Proposition 2:

Proof: Let's denote $Eq^{(c)*}$ as the expected child health in case c. The core solution shows that

$$Eq^{(c)*} = \begin{cases} \mu(1+\delta_1) + \lambda(1-\mu)(\delta_1+1) + \lambda^2[(1+\delta_2)\mu - (1-\mu)(1+\delta_1)] & \text{case 1} \\ \mu + \lambda(1-\mu) + \lambda^2(2\mu-1) + \lambda\tilde{\lambda}2(1-\mu) + \tilde{\lambda}^2[(\delta_1+\delta_2)\mu - 2(1-\mu)] & \text{case 2} \\ \frac{1}{2}\mu(1+\lambda)^2 + (1-\mu)(1+\lambda)\tilde{\lambda} + \tilde{\lambda}^2[(\delta_1+\delta_2)\mu - 2(1-\mu)] & \text{case 3} \end{cases}$$

We show that $Eq^{(1)*} > Eq^{(2)*} \geq Eq^{(3)*}$.

$$\begin{aligned} & Eq^{(1)*} - Eq^{(2)*} \\ = \; & \mu\delta_1 + \lambda(1-\mu)\delta_1 + \lambda^2(1+\delta_2)\mu - \lambda^2(1+\delta_1)(1-\mu) - \lambda^2(2\mu-1) \\ - \; & 2\lambda\tilde{\lambda}(1-\mu) - \tilde{\lambda}^2(\delta_1+\delta_2)\mu + 2\tilde{\lambda}^2(1-\mu) \\ = \; & \lambda(1-\mu)\delta_1 + \lambda^2[\delta_2\mu - (2+\delta_1)(1-\mu)] + 2\tilde{\lambda}(\tilde{\lambda}-\lambda)(1-\mu) + \delta_1\tilde{\lambda}(1-\lambda)^2 \\ > \; & 0 \end{aligned}$$

$$\begin{aligned} & Eq^{(2)*} - Eq^{(3)*} \\ = \; & \mu + \lambda(1-\mu) + \lambda^2(2\mu-1) + 2\lambda\tilde{\lambda}(1-\mu) - [\frac{1}{2}\mu(1+\lambda)^2 + (1-\mu)(1+\lambda)\tilde{\lambda}] \\ = \; & \frac{1}{2}\mu(1-\lambda)^2 - (1-\mu)(1-\lambda)(\tilde{\lambda}-\lambda) \\ \geq \; & \frac{1}{2}(1-\lambda)[\mu(1-\lambda) - 2(1-\mu)(\frac{1+\lambda}{2} - \lambda)] = \frac{1}{2}(1-\lambda)^2[\mu - (1-\mu)] \\ > \; & 0 \end{aligned}$$

Proof of Proposition 3:

Let's denote $EU^{(c)*}$ as the average equilibrium expected marital output in case c. We show that $EU^{(1)*} > EU^{(2)*} > EU^{(3)*}$.

Given the equilibrium matching and expected marital output for each type of couples, the total expected marital output in case 1 is as follows.

$$\begin{aligned} EU^{(1)*} &= \mu t^{rr} + (1-\mu)t^{rp} + (\delta_1\mu + \mu - 1)t^{pp} + (1+\delta_1)(1-\mu)\lambda t^{pr} \\ &+ [\mu - (1+\delta_1)(1-\mu)]\lambda^2 t^{rr} + (2+\delta_1)(1-\mu)\lambda^2 t^{rp} + [\delta_2\mu - (2+\delta_1)(1-\mu)]\lambda^2 t^{pp} \\ EU^{(2)*} &= [\mu + (1-\mu)\lambda + (2\mu-1)\lambda^2]t^{rr} + 2(1-\mu)\lambda\tilde{\lambda}t^{rp} + [(\delta_1+\delta_2)\mu - 2(1-\mu)]\tilde{\lambda}^2 t^{pp} \\ EU^{(3)*} &= \frac{1}{2}\mu(1+\lambda)^2 t^{rr} + (1-\mu)(1+\lambda)\tilde{\lambda}t^{rp} + [(\delta_1+\delta_2)\mu - 2(1-\mu)]\tilde{\lambda}^2 t^{pp} \end{aligned}$$

$$\begin{aligned} & EU^{(2)*} - EU^{(3)*} \\ = \ & [\mu + (1-\mu)\lambda + (2\mu-1)\lambda^2]t^{rr} + 2(1-\mu)\lambda\tilde{\lambda}t^{rp} - [\frac{1}{2}\mu(1+\lambda)^2 t^{rr} + (1-\mu)(1+\lambda)\tilde{\lambda}t^{rp}] \\ = \ & \frac{\mu}{2}[1-\lambda]t^{rr} + (1-\mu)[\lambda t^{rr} - \tilde{\lambda}t^{rp}] > 0 \end{aligned}$$

$$
\begin{aligned}
& EU^{(1)*} - EU^{(2)*} \\
= \;& (1-\mu)(t^{rp} - \lambda t^{rr}) - \delta_1(1-\mu)\lambda^2 t^{rr} + (\delta_1\mu + \mu - 1)t^{pp} + (1+\delta_1)(1-\mu)\lambda t^{rp} \\
& +(2+\delta_1)(1-\mu)\lambda^2 \Delta t^{rp}_{pp} - 2(1-\mu)\tilde{\lambda}[\lambda t^{rp} - \tilde{\lambda} t^{pp}] + [\delta_2\mu\lambda^2 - (\delta_1+\delta_2)\mu\tilde{\lambda}^2]t^{pp} \\
= \;& (1-\mu)[t^{rp} - \lambda t^{rr} - (t^{pp} - \lambda t^{rp}) + \delta_1\lambda(t^{rp} - \lambda t^{rr}) + (2+\delta_1)\lambda^2\Delta t^{rp}_{pp} \\
& -2\tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda} t^{pp}) + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
> \;& (1-\mu)[\delta_1\lambda(t^{rp} - \lambda t^{rr}) + (2+\delta_1)\lambda^2\Delta t^{rp}_{pp} - 2\tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda} t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
> \;& (1-\mu)[\delta_1\lambda(1-\lambda)t^{rp} - \delta_1\lambda\Delta t^{rp}_{pp} + (2+\delta_1)\lambda^2\Delta t^{rp}_{pp} - 2\tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda} t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
= \;& (1-\mu)[\delta_1\lambda(1-\lambda)t^{pp} - 2(\tilde{\lambda} - \lambda)(\lambda t^{rp} - \tilde{\lambda} t^{pp} + \lambda t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
> \;& (1-\mu)[\delta_1\lambda(1-\lambda)t^{pp} - (1-\lambda)(\lambda t^{rp} - \tilde{\lambda} t^{pp} + \lambda t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
= \;& (1-\mu)[(\delta_1 - 1)\lambda(1-\lambda)t^{pp} - (1-\lambda)(\lambda t^{rp} - \tilde{\lambda} t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
> \;& (1-\mu)[-(1-\lambda)(\lambda t^{rp} - \tilde{\lambda} t^{pp})] + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} \\
> \;& -(1-\mu)(1-\lambda)^2 t^{pp} + \mu\delta_1\tilde{\delta_2}(1-\lambda)^2 t^{pp} = (1-\lambda)^2 t^{pp}[\mu\delta_1\tilde{\delta_2} - (1-\mu)] > 0
\end{aligned}
$$

because of assumption 1, and the assumptions $\delta_1 \geq 1, \tilde{\delta_2} \geq \frac{1}{2}$.

Proof of Proposition 4:

Proof of Proposition 4(1):

Let $\zeta_g^{(c)*}$ denote the difference in average expected utility between type 1 and type 0 individuals of gender g in case c, where $g \in \{W, M\}; c \in \{1, 2, 3\}$.

$$\begin{aligned}
\zeta_M^{(1)*} &= \frac{1}{1+\delta_1}\Delta t_{pp}^{rp} + \lambda t^{rp} - \lambda^2(t^{rr} - \Delta t_{pp}^{pr}) - \frac{1}{1+\delta_2}\lambda^2 \Delta t_{pp}^{pr} \\
\zeta_W^{(1)*} &= \frac{1}{1+\delta_1}(t^{rr} - \Delta t_{pp}^{rp}) - \lambda t^{rp} + \lambda^2(t^{rr} - \Delta t_{pp}^{pr}) + \frac{\delta_1}{1+\delta_1}t^{pp} - \frac{1}{1+\delta_2}[\lambda^2(t^{rr} - \Delta t_{pp}^{pr}) + \delta_2\lambda^2 t^{pp}] \\
\zeta^{(1)*} &= \zeta_M^{(1)*} + \zeta_W^{(1)*} \\
&= \frac{1}{1+\delta_1}(t^{rr} + \delta_1 t^{pp}) - \frac{1}{1+\delta_2}[\lambda^2 t^{rr} + \delta_2\lambda^2 t^{pp}]
\end{aligned}$$

$$\begin{aligned}
\zeta_M^{(2)*} &= \frac{1}{1+\delta_1}[\lambda(1-\lambda)t^{rr} + \tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda}t^{pp})] - \frac{1}{1+\delta_2}\tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda}t^{pp}) \\
\zeta_W^{(2)*} &= \frac{1}{1+\delta_1}[(1-\lambda+\lambda^2)t^{rr} - \tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda}t^{pp}) + \delta_1\tilde{\lambda}t^{pp}] - \frac{1}{1+\delta_2}[\lambda^2 t^{rr} - \tilde{\lambda}(\lambda t^{rp} - \tilde{\lambda}t^{pp}) + \delta_2\lambda\tilde{\lambda}t^{pp}] \\
\zeta^{(2)*} &= \zeta_M^{(2)*} + \zeta_W^{(2)*} \\
&= \frac{1}{1+\delta_1}(t^{rr} + \delta_1\tilde{\lambda}t^{pp}) - \frac{1}{1+\delta_2}[\lambda^2 t^{rr} + \delta_2\lambda\tilde{\lambda}t^{pp}]
\end{aligned}$$

$$\begin{aligned}
\zeta_M^{(3)*} &= \frac{1}{1+\delta_1}\tilde{\lambda}(t^{rp} - \tilde{\lambda}t^{pp}) - \frac{1}{1+\delta_2}(\lambda\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp}) \\
\zeta_W^{(3)*} &= \frac{1}{1+\delta_1}[\frac{1+\lambda}{2}t^{rr} - (\frac{1+\lambda}{2}\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp}) + \delta_1\tilde{\lambda}t^{pp}] - \frac{1}{1+\delta_2}[\frac{1+\lambda}{2}\lambda t^{rr} - (\frac{1+\lambda}{2}\tilde{\lambda}t^{rp} - \tilde{\lambda}^2 t^{pp}) \\
&+ \delta_2\lambda\tilde{\lambda}t^{pp}] \\
\zeta^{(3)*} &= \zeta_M^{(3)*} + \zeta_W^{(3)*} \\
&= \frac{1}{1+\delta_1}(\frac{1+\lambda}{2}t^{rr} + \frac{1-\lambda}{2}\tilde{\lambda}t^{rp} + \delta_1\tilde{\lambda}t^{pp}) - \frac{1}{1+\delta_2}[\frac{1+\lambda}{2}\lambda t^{rr} - \frac{1-\lambda}{2}\tilde{\lambda}t^{rp} + \delta_2\lambda\tilde{\lambda}t^{pp}]
\end{aligned}$$

Thus,

$$\begin{aligned}
\zeta^{(1)*} - \zeta^{(2)*} &= \frac{\delta_1}{1+\delta_1}(1-\tilde{\lambda})t^{pp} - \frac{\delta_2}{1+\delta_2}(\lambda^2 - \lambda\tilde{\lambda})t^{pp} \\
&\propto \delta_1(1-\tilde{\lambda}) + \delta_2\lambda(\tilde{\lambda}-\lambda) > 0 \\
\zeta^{(2)*} - \zeta^{(3)*} &= \frac{1}{1+\delta_1}\frac{1-\lambda}{2}[t^{rr} - \tilde{\lambda}t^{rp}] - \frac{1}{1+\delta_2}\frac{1-\lambda}{2}[\tilde{\lambda}t^{rp} - \lambda t^{rr}] \\
&\propto \frac{1}{1+\delta_1}[t^{rr} - \tilde{\lambda}t^{rp}] - \frac{1}{1+\delta_2}[\tilde{\lambda}t^{rp} - \lambda t^{rr}] \\
&> \frac{1}{1+\delta_1}[t^{rr} - \tilde{\lambda}t^{rp}] - \frac{1}{1+\delta_1}[\tilde{\lambda}t^{rp} - \lambda t^{rr}] \\
&\propto (1+\lambda)t^{rr} - 2\tilde{\lambda}t^{rp} \geq (1+\lambda)(t^{rr} - t^{rp}) > 0
\end{aligned}$$

Therefore, the welfare difference between type 1 and type 0 shows

$$\zeta^{(1)*} > \zeta^{(2)*} > \zeta^{(3)*}$$

Proof of Proposition 4(2):

Let $\xi_g^{(c)*}$ denote the difference in average expected utility between type r and type p individuals of gender g in case c, where$g \in \{W, M\}; c \in \{1, 2, 3\}$. It is easy to show that, for gender g, $g \in \{W, M\}$, $\zeta^{(1)*} < \zeta^{(2)*}$ but $\zeta^{(2)*} > \zeta^{(3)*}$.

$$\begin{aligned}
\xi_M^{(1)*} - \xi_M^{(2)*} &= \frac{1}{2}(1+\lambda^2)\Delta t_{pp}^{rp} + (\frac{1}{2} - \tilde{\delta}_1)[\lambda t^{rp} - \lambda^2(\Delta t_{rp}^{rr} + t^{pp})] \\
&- \frac{1}{2}\lambda(1-\lambda)t^{rr} - \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) \\
\xi_W^{(1)*} - \xi_W^{(2)*} &= \frac{1}{2}[(1+\lambda^2)(t^{rr} - \Delta t_{pp}^{rp})] - (\frac{1}{2} - \tilde{\delta}_1)[\lambda t^{rp} - \lambda^2(\Delta t_{rp}^{rr} + t^{pp})] - (\tilde{\delta}_1 + \tilde{\delta}_2\lambda^2)t^{pp} \\
&- \frac{1}{2}(1-\lambda+2\lambda^2)t^{rr} + \tilde{\lambda}(\lambda t^{pr} - \tilde{\lambda} t^{pp}) + \tilde{\lambda}^2 t^{pp} \\
\xi^{(1)*} - \xi^{(2)*} &= \xi_M^{(1)*} - \xi_M^{(2)*} + \xi_W^{(1)*} - \xi_W^{(2)*} \\
&= \tilde{\lambda}^2 t^{pp} - (\tilde{\delta}_1 + \tilde{\delta}_2\lambda^2)t^{pp} = -\tilde{\delta}_1\tilde{\delta}_2(1-\lambda)^2 t^{pp} < 0
\end{aligned}$$

$$\begin{aligned}
\xi_M^{(2)*} - \xi_M^{(3)*} &= \frac{1}{2}\lambda(1-\lambda)t^{rr} - \frac{1-\lambda}{2}\tilde{\lambda} t^{pr} \\
\xi_W^{(2)*} - \xi_W^{(3)*} &= \frac{1}{2}[1-\lambda+2\lambda^2 - \frac{(1+\lambda)^2}{2}]t^{rr} + \frac{1-\lambda}{2}\tilde{\lambda} t^{pr} \\
\xi^{(2)*} - \xi^{(3)*} &= \xi_M^{(2)*} - \xi_M^{(3)*} + \xi_W^{(2)*} - \xi_W^{(3)*} \\
&= \frac{(1-\lambda)^2}{4} t^{rr} > 0
\end{aligned}$$

B Complementary Empirical Analysis

B.1 Complementarity of Parental Health on Producing Children Health

We provide suggestive evidence for the complementarity of parental health on producing children health.

The data on spousal health before age 15 was obtained from the 2014 Life History Survey of the China Health and Retirement Longitudinal Study (CHARLS), which collects information on individuals aged 45 and above and their families. We do not use CFPS data here as it does not contain information on the premarital health status. To minimize reporting errors, we only included respondents born after 1930. All the spouses in our sample were married before 2003 (which avoids the potential effects of the PHE repeal policy). We use two variables to measure parents' health. One is the dummy variable of "having been bedridden/hospitalized before age 15" based on the questions "Before age 15, have you ever been bedridden for a month or longer due to health reasons?" and "Before age 15, have you ever been hospitalized for a month or longer due to health reasons?" The dummy variable equals to 1 if the answer to either question is "yes", and 0 otherwise. The other measure is the dummy variable for "being healthy before age 15" based on the question "Before age 15, would you say that compared to other children of the same age, you were: 1. much healthier; 2. somewhat healthier; 3. about average; 4. somewhat less healthy; 5. much less healthy". We set the dummy variable to 1 if the answer was 3 or below, and 0 otherwise.

Besides, we constructed two variables measuring children health. The information is also drawn from the 2014 Life History Survey of the CHARLS. One is a dummy variable for "healthy birth" based on the question "Did this pregnancy end up with induced abortion, miscarriage/natural abortion, or stillbirth?" For induced abortion, we further restricted it to cases where the reason was that "the child is not healthy according to the ultrasound result." If none of the pregnancies of a couple ended up with these cases, we set the dummy variable for healthy birth to 1; otherwise, we set it to 0. Another measure is the dummy variable for parents' reported child health. This variable was constructed using the question "How is this child's health? 1. Very good, 2. Good, 3. Fair, 4. Poor, 5. Very poor?" We set the dummy variable to 1 if the answer was 3 or below, and 0 otherwise.

Panel A in Table B1a shows mean of "health birth" for three groups of couples. Columns (1), (2), and (3) are for the couples with 2, 1, and 0 spouse having been bedridden or hospitalized before age 15, respectively. The couples become healthier from Column (1) to (3). Columns (4), (5), and (6) are for couples with 0, 1, and 2 spouses reporting being healthy before age 15. The couples becomes healthier from Column (4) to (6). The mean of "Healthy birth" is 0.875, 0.881, and 0.894 in Columns (1) to (3), respectively, and the mean is 0.878, 0.879, and 0.897 in Columns (4) to (6). Panel B shows the difference of the mean of "health birth" between Columns (2) and (1), (3) and (2), (5) and (4), and (6) and (5). We see that the difference between Columns (3) and (2) is larger than that between Columns (2) and (1). We also see that the difference between Columns (6) and (5) is larger than that between Columns (5) and (4). Table B1b shows the similar information as in Table B1a but focuses on "children health reported by parents".

Results in both Table B1a and Table B1b provide suggestive evidence for the complementarity of parental health on producing children health.

Table B1a. Complementarity of Spousal Premarital Health on Child Health Production: Health Birth

	Healthy birth					
	Number of spouses having been bedridden/hospitalized before age 15			Number of spouses reporting being healthy before age 15		
	2	1	0	0	1	2
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A						
Mean	0.875	0.881	0.894	0.878	0.879	0.897
S.D.	(0.336)	(0.324)	(0.308)	(0.328)	(0.327)	(0.305)
Observations	32	589	4,657	115	1,046	4,117
Panel B						
	Between group difference			Between group difference		
		(2)-(1)	(3)-(2)		(5)-(4)	(6)-(5)
Difference		0.006	0.013		0.000	0.018*
S.E.		[0.059]	[0.014]		[0.032]	[0.011]

Notes: (1) This table is based on couple-level observations. We constructed a dummy variable for "healthy birth" based on the question "Did this pregnancy end up with induced abortion, miscarriage/natural abortion, or stillbirth?" For induced abortion, we further restricted it to cases where the reason was that "the child is not healthy according to the ultrasound result." If none of the pregnancies of a couple ended up with these cases, we set the dummy variable for healthy birth to 1; otherwise, we set it to 0. (2) Standard deviations are reported in parentheses; standard errors are reported in square brackets. *, **, *** denote statistical significance at the 10%, 5%, and 1% levels.

Table B1b. Complementarity of Spousal Premarital Health on Child Health Production: Children Health Reported by Parents

	Children health reported by parents					
	Number of spouses having been bedridden/hospitalized before age 15			Number of spouses reporting being healthy before age 15		
	2	1	0	0	1	2
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A						
Mean	0.909	0.915	0.956	0.840	0.935	0.959
S.D.	(0.290)	(0.280)	(0.205)	(0.368)	(0.247)	(0.199)
Observations	66	1,230	9,889	268	2,213	8,704
Panel B						
	Between group difference			Between group difference		
		(2)-(1)	(3)-(2)		(5)-(4)	(6)-(5)
Difference		0.006	0.041***		0.095***	0.024***
S.E.		[0.035]	[0.006]		[0.017]	[0.005]

Notes: (1) This table is based on children-level observations. The dummy variable for parents' reported child health was constructed using the question "How is this child's health? 1. Very good, 2. Good, 3. Fair, 4. Poor, 5. Very poor?" We set the dummy variable to 1 if the answer was 3 or below. (2) Standard deviations are reported in parentheses; standard errors are reported in square brackets. *, **, *** denote statistical significance at the 10%, 5%, and 1% levels.

Table B2. Correlation between the PHE and Other Provincial Variables in 2002

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Log(GDP per capita)	0.229***								0.234
	(0.035)								(0.342)
Log(consumption per capita)		0.276***							-0.005
		(0.042)							(0.363)
Share of primary industry in GDP			-1.456***						-0.369
			(0.298)						(1.074)
Num. of health institutes per 10,000 people				0.012					0.008
				(0.029)					(0.023)
Num. of doctors per 10,000 people					0.014***				0.006
					(0.002)				(0.008)
Num. of civil affair staffs per 10,000 people						-0.007			0.035
						(0.139)			(0.155)
Log(government health expenditures per capita)							0.172***		-0.103
							(0.039)		(0.129)
Infectious diseases (categories A and B) infection rate								-0.001	0.003
								(0.005)	(0.006)
Observations	30	30	30	30	30	30	30	30	30
R-squared	0.466	0.447	0.316	0.004	0.280	0.000	0.255	0.002	0.503
P-value of F test									0.000

Notes: Robust standard errors in parentheses. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table B3. Balance Test Using 2005 Population Census

	(1) Education	(2) Number of siblings	(3) Ethnic minority
PHE rate×Post2003	0.073	-0.327	-0.042
	(0.080)	(0.187)	(0.029)
Observations	82,385	82,385	82,385
R-squared	0.168	0.097	0.474

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors are reported in parentheses, two-way clustered by province and year of marriage; cohorts married within two years before/after the policy shock are used.

C Robustness Checks

C.1 Alternative Control Variables

We show in the main text that predetermined individual characteristics are balanced across provinces and in the pre- and postrepeal periods. Lee and Lemieux (2010) suggest that if the distribution of these predetermined characteristics across provinces around the policy shock is balanced, the inclusion of these controls should have little effect on the baseline estimates. Results to the contrary may suggest that our research design is invalid. As a robustness check, we include these individual-level predetermined variables in the baseline regression. The estimation results are reported in Panel A of Table C1. Our key estimates remain robust in both magnitude and statistical significance.

To investigate whether our baseline findings are driven by the choice of fixed effects, we substitute county fixed effects with province fixed effects in Panel B of Table C1. We still find highly robust estimates. We also interact the province-level predetermined variables with a set of year-of-marriage dummies instead of a post 2003 dummy to enable more flexible time-varying effect on SWB. The results in Panel C of Table C1 confirm our estimates' robustness in magnitude, although the statistical significance decreases to some extent, possibly due to the increased risk of model overfitting.

C.2 Inclusion of Individuals Married in 2003

In the main analysis, we drop individuals married in 2003 to ensure clear identification of the treatment and control groups. To investigate how the inclusion of individuals married in 2003 changes the estimates, we include them in the sample and define their treatment status according to their reported month of marriage. Individuals are classed in the policy affected group if their marriage month is after October 2003 and in the unaffected group otherwise. Given that individuals registering for marriage right after October might have already conducted the PHE due to the time gap between the PHE and the marriage registration, including individuals married in 2003 could cause a downward bias. The results in

Table C2 show that the coefficients remain robust.

C.3 Using Emotional Well-Being Measures as Alternative Outcomes

While the evaluative well-being measure is a powerful indicator of an individual's overall utility and is widely used in the literature studying SWB, emotional well-being captures the actual experience in a recent period (Kahneman and Deaton, 2010). We therefore use emotional well-being as a complement to measure SWB. We obtain three specific measures from the CFPS survey: the frequencies of feeling depressed, feeling nervous, and feeling restless. Their value ranges from 1 (corresponding to never in the last month) to 5 (almost every day in the last month). The regression results in Table C3 show that the PHE repeal resulted in a significant increase in the frequency of depression, nervousness, and restlessness. This confirms that our conclusion holds when different measurements of SWB are used.

C.4 Using the Urban Sample

In the main analysis, we focus on individuals with a rural *hukou*. As a comparison, we investigate the impact of the PHE repeal on individuals with an urban *hukou*. Due to *hukou* restrictions, the marriage markets in rural and urban areas are segmented (Han and Shi, 2019). The empirical results are shown in Table C4. No coefficients are significant, showing that the repeal of the PHE had no significant impacts on SWB for urban people. One possible reason is that the prevalence of infectious disease in urban areas is lower than that in rural areas (Wang et al., 2019). Compared with rural people, urban people have better access to prenatal care, which may serve as a substitute for the role of the PHE for healthy childbearing.

C.5 Using Different Ways of Calculating Standard Errors

We determine whether the baseline statistical inference is robust to the choice of methods to calculate standard errors. We experiment with two alternative clustering methods. One is to calculate the standard errors by clustering over provinces, allowing the error terms to be correlated across individuals in the same province. Considering the small number of clustering units, 25 in this case, we implement the wild bootstrap procedure (Cameron et al., 2008). The other practice is to calculate standard errors by clustering over counties, allowing the error terms to be correlated within the same county. There are 192 counties in our sample. The results shown in Table C5 are robust.

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Table C1. Using Alternative Control Variables

	(1) Life Satisfaction	(2) Family Satisfaction	(3) Happiness
Panel A: Inclusion of predetermined personal characteristics			
PHE rate×Post2003	-1.705***	-1.375***	-5.571***
	(0.383)	(0.355)	(0.742)
Observations	2,316	2,316	2,316
R-squared	0.144	0.152	0.156
Predetermined controls	Yes	Yes	Yes
Panel B: Province FE in place of county FE			
PHE rate×Post2003	-0.915***	-0.811**	-3.211***
	(0.137)	(0.296)	(0.715)
Observations	3,213	3,213	3,213
R-squared	0.042	0.041	0.070
Panel C: Provincial predetermined variables interacted with year-of-marriage dummies			
PHE rate×Post2003	-1.081**	-0.852	-3.432***
	(0.477)	(0.650)	(0.960)
Observations	3,213	3,213	3,213
R-squared	0.151	0.152	0.163

Notes: We include provincial controls and year-of-marriage fixed effects in all regressions. Panel A and C include county fixed effects, while Panel B includes province fixed effects. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively. Predetermined controls include the respondent's father's education, mother's education, own education, number of siblings, ethnic minority, and weeks not living with parents before age 12 (taking logarithm).

Table C2. Including Individuals Married in 2003

	(1)	(2)	(3)
	Life Satisfaction	Family Satisfaction	Happiness
PHE rate×Post2003	-0.978***	-0.901***	-2.681***
	(0.306)	(0.282)	(0.428)
Observations	3,445	3,445	3,445
R-squared	0.119	0.119	0.126

Notes: We include provincial controls and county and year-of-marriage fixed effects in all regressions. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table C3. Emotional Well-being as Outcome Variables

	(1)	(2)	(3)
	Depression	Nervousness	Restlessness
PHE rate×Post2003	0.939**	0.838***	0.903***
	(0.338)	(0.160)	(0.228)
Observations	3,210	3,212	3,212
R-squared	0.119	0.125	0.117
Mean of y	1.798	1.629	1.543

Notes: We include provincial controls and county and year-of-marriage fixed effects in all regressions. The standard errors are reported in parentheses, two-way clustered by province and year of marriage. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table C4. Using Urban Sample

	(1)	(2)	(3)
	Life Satisfaction	Family Satisfaction	Happiness
PHE rate×Post2003	-0.552	-0.195	-0.731
	(0.909)	(0.908)	(0.469)
Observations	1,295	1,295	1,293
R-squared	0.190	0.186	0.161
Mean of y	3.763	3.894	7.742

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The standard errors two-way clustered by province and year of marriage are reported in parentheses. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.

Table C5. Different Ways of Clustering

	(1)	(2)	(3)
	Life Satisfaction	Family Satisfaction	Happiness
PHE rate×Post2003	-1.076	-0.849	-3.412
	[0.02]**	[0.04]**	[0.01]**
	{0.541}**	{0.471}*	{0.471}***
Observations	3,213	3,213	3,213
R-squared	0.122	0.125	0.132

Notes: We include provincial controls and county and year-of-marriage fixed effects in all columns. The p-value with standard errors clustered by province using the wild bootstrap procedure are reported in square brackets. The standard errors clustered by county are reported in curly brackets. ***, **, and * denote statistical significance at 1%, 5%, and 10%, respectively.