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The Dark Side of Competition

Gender Differences

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ABSTRACT

The literature has placed great emphasis on the advantages of competition on market efficiency while ignoring the downside of competition on health. Using a natural experiment in Taiwan, we show that excessive competition comes at a health cost. In the late 1940s, half a million soldiers retreated to Taiwan from Mainland China after a civil war. They were initially not allowed to get married until the marriage ban was essentially lifted in 1959. As a large number of soldiers flooded the marriage market, men faced much stronger mating competition than before, which in turn increased the likelihood of male depression and mortality.

Keywords: competition, sex ratio, mortality, health outcomes, gender

JEL Codes: I18, J12

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1. INTRODUCTION

A burgeoning body of literature has examined why women are less competitive than men (Gneezy, Niederle, and Rustichini 2003; Niederle and Vesterlund 2011). Women's relatively lower level of competitiveness has been regarded by some researchers as a key cause of their worse labor market outcomes compared to their male counterparts. This body of literature implicitly assumes that more competitiveness is associated with better outcomes. However, increasing competition may come at a health cost.

Competition results in stress. From a biological perspective, under stressful conditions, human brain mobilizes mechanisms that produce energy for coping and suppresses other mechanisms that store energy or use it for growth, repair, and surveillance against pathogens. Part of this adaptive response is to raise blood pressure along with a variety of biochemical changes. When the stress is chronic, the high blood pressure in interaction with stress-induced biochemical changes could cause damages that lead to life-threatening diseases such as coronary heart disease, stroke, and kidney disease (Sterling and Eyer 1981).

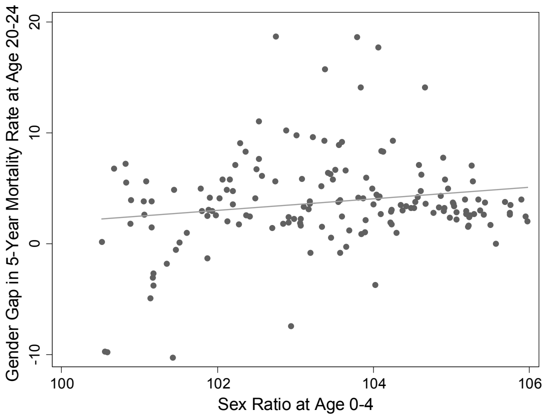
In another literature of sociological stress research, stressful experiences have been shown to harm both physical and mental health (see Thoits 2010 for a recent review). In particular, one major finding from this literature is that differential exposure to stressful experiences produces health inequalities within a population. It has been widely observed that men on average live shorter lives than women and suffer higher mortality rates at almost all ages probably because men generally experience more competition stress than women through their life course. It is likely that excessive competition may harm men's health, resulting in a higher mortality rate compared to women in the long run.

Because many factors, including biological, cultural, and environmental factors, can affect the gender difference in health outcomes and mortality rates (Yin 2007), it is hard to separate the long-term impact of exposure to over-competition from other factors without using an experimental approach. However, it is unethical to apply laboratory or field experiments to human subjects to study the health impacts of over-competition. Moreover, it is a great challenge to measure the degree of competition as people face many types of competition in daily life.

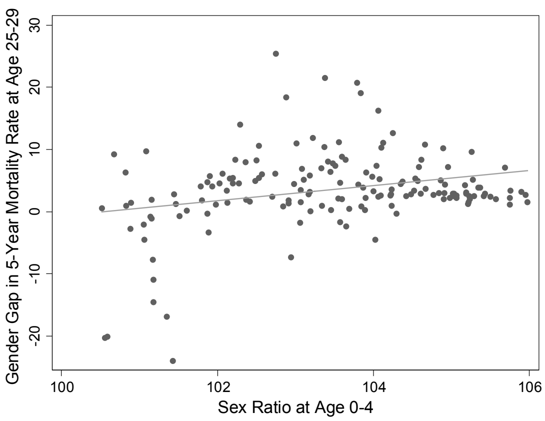
In this paper, we focus on mating competition, the intensity of which, to a large extent, can be accurately measured by the sex ratio among people at marriageable age. Mating is an integral part of human life. An increase in sex ratio (more men relative to women) will intensify male competition but reduce women's competitive pressure in the marriage market. In this paper we examine the impact of mating competition, measured by sex ratio, on the gender difference in mortality rate and other health outcomes.

We first use cross-country data to see if sex ratio is associated with a gender difference in mortality rate. Figure 1.1 plots the gender difference in five-year mortality rates across 160 countries and territories for cohorts who were 0 to 4 years old in 1980 and were later observed when they reached 20 to 24, 25 to 29, and 30 to 34 years old, respectively. The sex ratio for age group 0 to 4 in 1980 to some extent measures the degree of mating competition when the cohort enters the marriage market later on. It is apparent from the figure that the gender gap in adult mortality rate among the prime marriage-age cohort is significantly higher in countries with more skewed sex ratios at birth than those with more balanced sex ratios at birth, indicating that an excess number of males relative to females at birth may harm male mortality in later life.

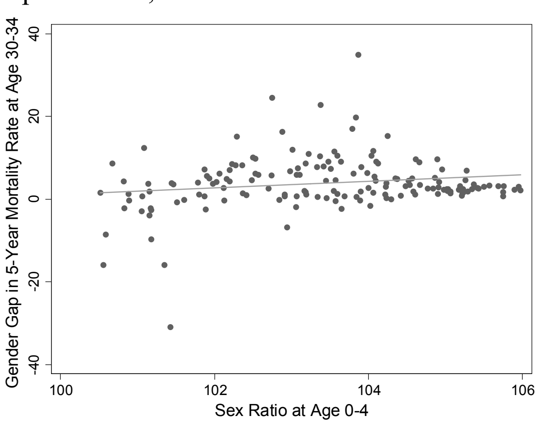
Figure 1.1 Five-year adult mortality rates and sex ratios at birth across 160 countries



Slope = 0.523; *t* = 2.29.



Slope = 1.206; *t* = 3.52.



Slope = 0.788; *t* = 2.21.

Source: World Population Prospects, 2015 Revision (UN ESA-PD 2015).

Note: Sex ratio for the birth cohorts of age 0 to 4 in 1980. Countries at the top and bottom 10 percent of the sex ratio distribution are dropped. The final sample includes 160 countries and territories. Five-year mortality rates for age 20 to 24, 25 to 29, and 30 to 34 are measured as number of deaths per 1,000 people observed between 2000 and 2005, between 2005 and 2010, and between 2010 and 2015.

Of course, the bivariate plot shows just a correlation and does not imply any causal relationship. Gender differences in mortality rate are affected by many factors, such as biological, cultural, institutional, and personal behavioral factors. Unfortunately, cultural and institutional factors, which shape men's and women's attitudes toward competition (Gneezy, Leonard, and List 2009), can hardly be controlled for in cross-country analyses due to a lack of systematic data for all countries. Moreover, the use of aggregate data masks the important role of individual risk factors, such as smoking behavior (Case and Paxson 2005), on mortality. So the positive correlation between sex ratio at birth and later gender difference in mortality is at most suggestive.

In this study, we take advantage of a powerful natural experiment in Taiwan to study the long-term impact of mating competition on gender difference in health outcomes, including mortality. In the late 1940s, about half a million of Chiang Kai-shek's soldiers, mostly unmarried young men, retreated to Taiwan after a civil war against Mao Zedong's People's Liberation Army (PLA) in Mainland China, while the local Taiwanese civilian population was only 6 million at the time. Upon their arrival in Taiwan, Chiang's soldiers were immediately deployed in every county throughout the island to defend against the PLA's attacks from the west side of the Taiwan Strait.

For a long period of time, in determination of recovering the mainland, these soldiers were confined to military compounds and were not allowed to get married. The marriage ban was essentially lifted in 1959, injecting a large number of bachelors into the marriage market in Taiwan. Much to their dismay, local Taiwanese men at marriageable ages suddenly faced much stronger mating competition than before. Stress directly resulted from the enhanced mating competition may have cast a toll on the health of the men seeking a mate in the local marriage market. Indirectly, stress could also arise because these men had to work longer hours or start their own business in order to accumulate more wealth so that their competitiveness was enhanced in the marriage market.

The large-scale natural experiment in Taiwan—under a rather homogeneous culture and institution across counties—results in both geographical and temporal variations in the sex ratio in local marriage markets. The intensity of marriage market competition varied across counties mainly due to variation in the military deployment in each county. The exogenous variations in the sex ratio enable us to identify the long-term health impact of stress arising from mating competition. We track the mortality rate of the affected cohort and other cohorts in the past six decades and find that the enhanced male mating competition following the removal of the marriage ban elevates the likelihood of depression and death for men but not for women.

Our paper is related to the literature on gender differences in competiveness and several other strands of the literature. First, it relates to the literature on the relationship between stress and health. Price et al. (1994) and Gilbert et al. (2009) have shown that stress is a major risk factor for both physical and mental health problems. Our paper reveals that the particular form of stress stemming from elevated mating competition can have a negative health consequence, even a deadly one. Second, our paper contributes to the emerging body of literature on the consequences of sex ratio imbalances, see, for example, Wei and Zhang (2011a) on savings; Wei and Zhang (2011b) and Chang and Zhang (2015) on entrepreneurship; Wei, Zhang, and Liu (2012) on housing; and Edlund et al. (2013) on crime. One key difference of this paper from previous studies lies in its focus on the long-term health impacts of a skewed sex ratio.

Sex ratio imbalances have become a serious problem in many Asian countries, such as China and India. For example, in China, since the one-child policy (OCP) was introduced in 1979, due to a combination of factors (OCP, son-preference culture, and availability of ultrasound gender-identification technology (Bulte, Heerink, and Zhang 2011; Li, Yi, and Zhang 2011), sex ratios have increased from a natural rate of 106 in the 1970s to 120 in 2010, implying that six men compete for five women in the marriage age. Because the cohort born after OCP in Mainland China is still young, it is too early to discern the long-term health impact, if any. The findings in our paper foretell some looming health consequences for men born after OCP in Mainland China in the upcoming two or three decades.

Competition is a part of human nature. Overall, the economics literature has placed great emphasis on the positive side of competition on market efficiency while ignoring the downside of competition on health in the long run. Our paper provides empirical evidence that excessive competition comes at a cost. The dark side of competition should not be ignored.

The paper is arranged as follows. The next section introduces the background of the marriage policy in Taiwan and datasets used in the analyses. Section 3 presents the statistical evidence on the impact of mating competition on health outcomes for men and women. Section 4 concludes.

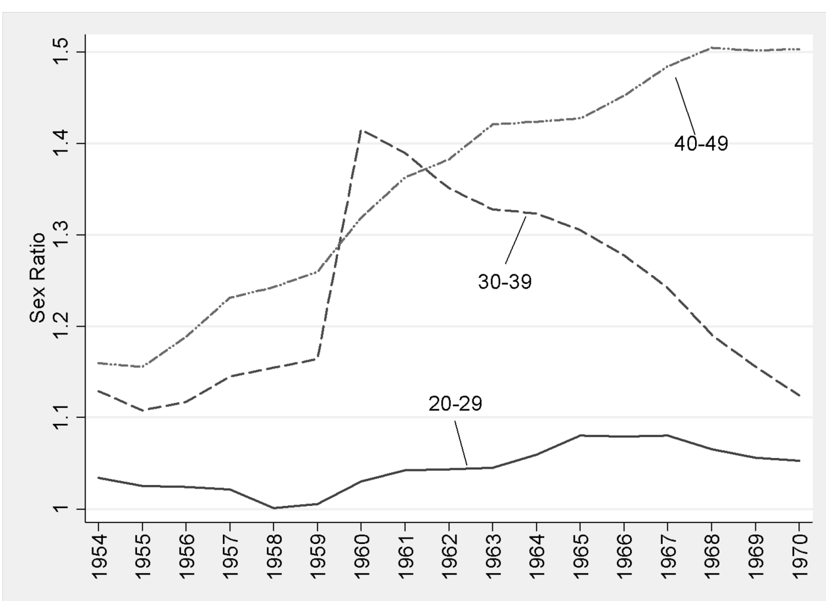
2. BACKGROUND AND DATA

In China, the Chinese Nationalist Party (also known as Kuomintang) led by Chiang Kai-shek and the Chinese Communist Party (CCP) led by Mao Zedong fought against each other in a civil war between 1945 and 1949. With the defeat of Kuomintang, a great influx of immigrants, including about half a million young unmarried male soldiers, retreated to Taiwan in the late 1940s (Barclay 1954; Jacoby 1967; Ho 1978; Chen and Yeh 1982; Liu 1986; Lin 2002). In contrast, the local population numbered merely 6 million. In this group of civil war immigrants, men outnumbered women by four to one (Francis 2011). Upon their arrival in Taiwan, Chiang's soldiers were immediately deployed in every county throughout the island to defend against the CCP's attacks.

For a long period of time after their arrival, the soldiers were confined to their military bases and not allowed to get married in order to remain vigilant and ready for any potential warfare (Lin 2002). In 1952, the marriage ban was formally written into a law called the Military Marriage Ordinance (MMO). The MMO forbade most active military personnel from getting married, except for military officers and technician sergeants. However, in August 1959, the ban was relaxed for most soldiers, except for male soldiers younger than 25, female soldiers younger than 20, and all soldiers who had served fewer than three years. This relaxation essentially made the MMO nonbinding for most immigrant soldiers, who either were already older than 25 or had served more than three years by 1959. This meant that they could finally get married, roughly 10 years after their arrival in Taiwan. With a sudden injection of about half a million bachelors into the marriage market, the effective sex ratio (the number of men to women at the marriage age between 15 and 49) immediately became skewed.

We use the imputed sex ratios that have been used in Chang and Zhang (2015). Figure 2.1 displays the imputed sex ratios at the Taiwan level (mthe whole island of Taiwan but excluding the small isles such as Penghu, Quemoy and Matsu that are also under Chiang's control) for three age cohorts, 20 to 29, 30 to 39, and 40 to 49. The sex ratio for the 30 to 39 cohort shot up from about 1.16 in 1959 to 1.41 in 1960, after the marriage ban was removed. The sex ratio for the older, 40 to 49, cohort steadily increased and peaked at greater than 1.5 in 1969 and 1970. With such high sex ratios, men had to look for younger women to get married. As a result, the marriage age gap between men and women had widened from about one or two years in the mid-1950s to more than five years in 1970 (Chang and Zhang 2015).

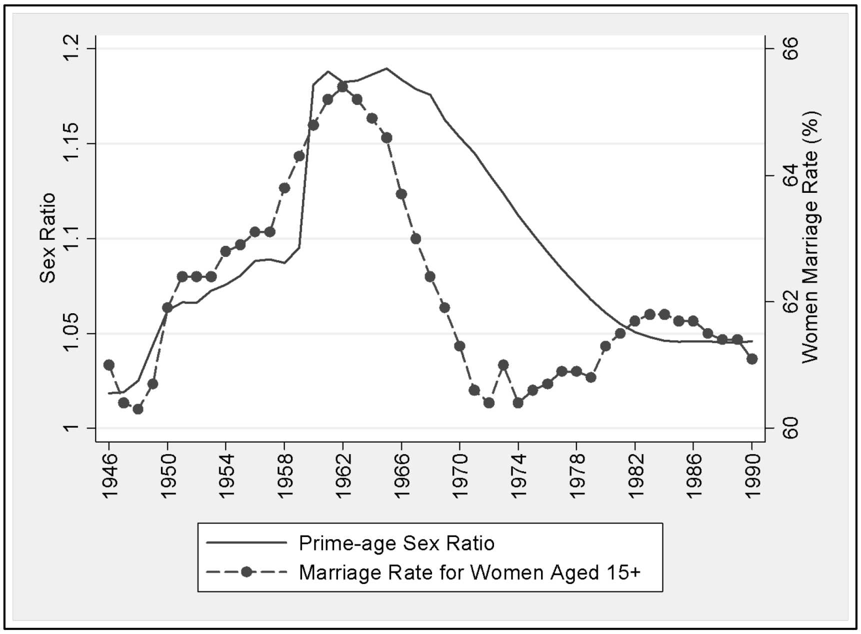
Figure 2.1 Imputed sex ratio by cohort



Source: Chang and Zhang (2015).

Figure 2.2 illustrates the imputed prime age sex ratio that covers age between 15 and 49 (solid line) at the Taiwan level. In our following analysis, we use the imputed prime age sex ratio at the county level as the key regressor. Overall, the prime age sex ratio rose from 1.02 in the mid-1940s to close to 1.2 in the 1960s. It then started to decline in the 1970s as the cohorts gradually aged out. Meanwhile, women's marriage rate (dotted line) increased from 62 percent in 1950 to 65 percent in 1962. As older men, mostly from Mainland China, increasingly dated and married younger women, young local Taiwanese men suddenly faced much greater competitive pressure in the marriage market.

Figure 2.2 Prime age (15–49) sex ratio and women's marriage rate



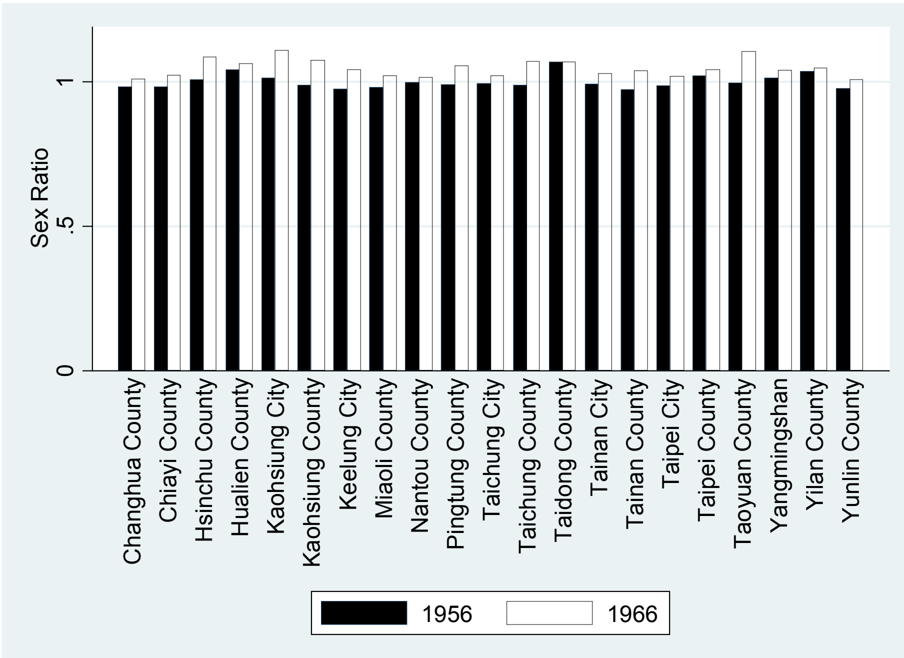
Source: Chang and Zhang (2015).

The imputed sex ratios at the Taiwan level do not separate the Mainland Chinese immigrants from the native Taiwanese due to the limitation of the official annual population data collected via household registration system. They also do not show county-level variations, which, as we mentioned before, resulted from the variations in the deployment of soldiers across counties. Figure 2.3 compares the county-level sex ratios for native Taiwanese (Panel A) and Mainland Chinese (Panel B) in 1956 and 1966 based on the Taiwan population census data, which include both civilians and military personnel. It can be seen that the sex ratios for native Taiwanese were rather balanced across counties and years, while the ratios for the Mainland Chinese immigrants display a great variation across counties and years. Such an asymmetric pattern between the sex ratios of these two groups suggests that internal migration in response to imbalanced sex ratio is less of a concern, because if male Mainland Chinese immigrants seeking a mate had an incentive to move to a county where the sex ratio was more balanced, the same would have had applied to Taiwanese men. Instead, the variation among Mainland Chinese immigrants across counties reflects more of the changes in military deployment.

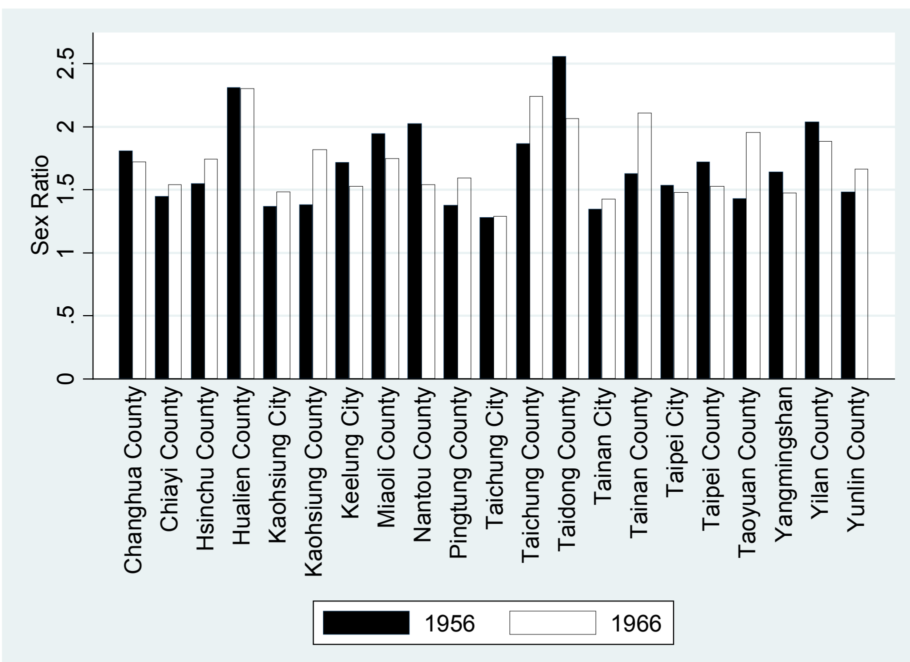
In this paper, we exploit the variations in sex ratios mainly resulted from the civil war immigration interacted with the marriage market policy change in 1959 to identify the impact of mating competition on men's and women's long-term health outcomes, including mortality, health status, depression, and life satisfaction. We use four datasets to capture the different outcome variables.

Figure 2.3 Sex Ratio of Native Taiwanese and Mainland Chinese in 1956 and 1966

Panel A: Native Taiwanese



Panel B: Mainland Chinese



Source: Chang and Zhang (2015).

Taiwan Manpower Survey 1978

The Taiwan Manpower Survey (TMS) reports labor market outcomes for a nationally representative sample of the civilian population aged 15 and older. TMS is administered by the Directorate-General of Budget, Accounting and Statistics in Taiwan. TMS collects, in addition to basic demographic information, detailed information about individual labor market outcomes, such as labor force participation, hours of work, and job characteristics. We use a subsample extracted from TMS in 1978, cohorts born between 1928 and 1958 who were aged 20 to 50 in 1978. We use this dataset to measure the impact of sex ratio imbalances on labor market outcome, such as work status and work hours.

National Survey of Living Conditions 1992–1994

The National Survey of Living Conditions (NSLC) is a national survey conducted by the Ministry of Interior to collect information about individual satisfaction with overall life and many other aspects of life, such as health, marriage, finance, education, and environment, along with basic demographic information. NSLC is a repeated cross-sectional survey of nationally representative samples of the civilian population aged 20 and older. We extract samples of cohorts born between 1926 and 1960 from the 1992, 1993, and 1994 waves of the NSLC. This dataset is used to study the impact of mating competition on life satisfaction.

Mortality Data 1981–2004

One key outcome variable is age-gender-specific mortality rate at the county level. The variable is defined as the ratio of number of deaths during aged 50 to 54 to total population in a specific gender-cohort. The number of deaths at the county level comes from the death registry for 1981 to 2004 from the Ministry of Health and Welfare. The population data are obtained from the 1980 Population and Housing Census in Taiwan. It was administered by the Ministry of the Interior. In this paper, we focus on the five-year mortality rate for the cohorts born between 1931 and 1950 when they were aged 50 to 54. See Appendix A for details about the calculation of the mortality rate. Since the death registry's information on the county of residence are no longer released to the public after 2004, we cannot observe the mortality rate for the cohort when they become older than 54. This limits our sample to the 50 to 54 age cohort.

The Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan 1996 and 2003

The Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan is a panel survey conducted by the Health Promotion Administration under the Ministry of Health and Welfare. The survey collects detailed information about demographics, health outcomes, employment history, residential history, economic conditions, and so forth. We use two samples drawn in 1996 and 2003, respectively. The 1996 sample consists of cohorts born between 1926 and 1946, and the 2003 sample includes cohorts born between 1946 and 1953. We use this dataset mainly to measure the impact on self-reported health status and depression.

As for the key explanatory variable, we employ the county-level effective sex ratio at age 20 as a measure of marriage competition. We choose age 20 as a time point when individuals entered the marriage market based on the average age of first marriage in the 1950s, which was 23 for men and 21 for women. Again, the sex ratio is defined as the ratio of men to women of age 15 to 49. The wide range of age is to reflect the fact that many soldiers from Mainland China had reached their 30s or 40s when the marriage ban was lifted in the 1959. As a robustness check, we also try various different definitions of sex ratio later in our empirical analysis.

3. EMPIRICAL RESULTS

Impact on Mortality Based on Aggregate Data

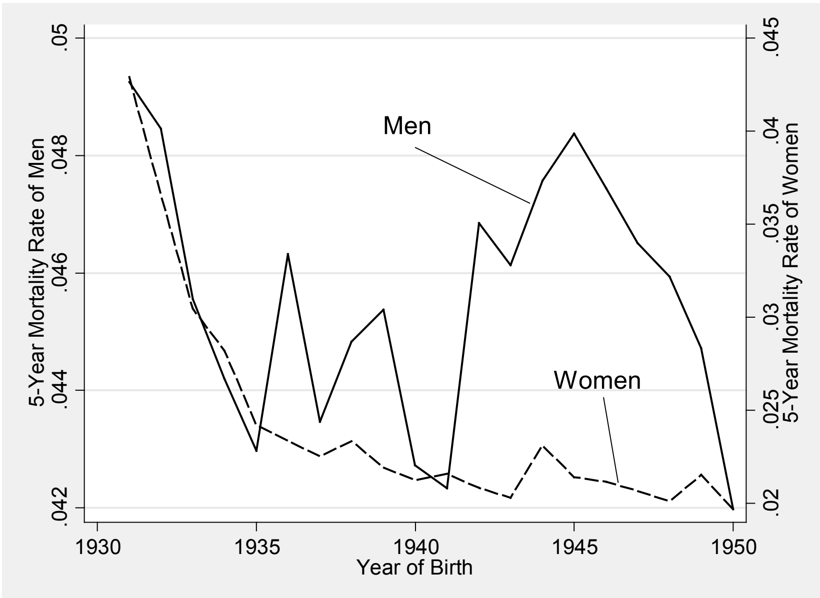
In this subsection, we provide some empirical evidence that exposure to a more competitive mating environment due to the removal of the marriage ban during their prime marriageable age increases men's likelihood of death in later life. Figure 3.1 displays the gender difference in mortality rate for age cohort 50 to 54 according to birth year. For those born before 1935, the gender difference in mortality rate is minimal. In 1959, they were at least 24 years old. Given that the average marriage age for men in 1959 was 22, most of them were supposedly married already. Therefore, the sudden tightening of the marriage market did not affect most men's marriage likelihood. This probably explains why men and women in this cohort showed little difference in mortality rate. By comparison, the male cohort born in the period between 1936 and 1950 was 9 to 23 years old in 1959 and was thereby likely subject to more intense mating competition than the older cohort when they entered the marriage market. As shown in the figure, the male mortality rate was much higher than the female mortality rate for this cohort, providing suggestive evidence that fiercer male mating competition in the marriageable age is associated with a higher likelihood of male mortality in later life. However, the positive correlation visualized in Figure 3.1 between elevated competitive pressure and a later higher male mortality rate may be driven by unobserved factors.

$$M_{cba}^{g} = S_{cba}\beta^{g} + D_{cb}\delta^{g} + \varepsilon_{cba}^{g}, \tag{1}$$

where M_cba^gis the county level mortality rate for males (g=m) or females (g=f) at age a for birth cohort b in county c, ε_cba^g is an error term, S_cba is the sex ratio when cohorts b was at age 20, and D_cb is a vector consisting of cohort and county dummies. Table 3.1 presents the regression results of five-year mortality rate at ages 50 to 54 for birth cohort 1931 to 1950 on the effective sex ratio at the county level that the cohort faced at age 20. The dependent variable is county-level sex-cohort-specific five-year mortality rates at age 50 to 54, which are defined as the ratio of the total deaths at ages 50 to 54 for each cohort to the cohort size at the end of age 49. Sex ratio S_cba is defined as the number of men per women for cohort 15 to 49. It is evaluated at the county level when the cohort reached age 20. It would be ideal to match the mortality data with actual residency at age 20, at which the sex ratio variable is evaluated. However, such information is not available. As a second best, we use the current residence in the 1980 population census as the proxy for residence at age 20 (panel A). In panel B, we use the residence in 1975 as another proxy for residence at age 20. In addition, all regressions include a full set of cohort dummies and county dummies to control for both cohort- and county-specific fixed effects. As shown in both panels, sex ratio at age 20 affects only men's mortality rate at ages 50 to 54, but not women's, and the sex difference is statistically significant. The sex ratio increased from slightly greater than 1.0 in the early 1950s to around 1.2 in 1960. Based on the estimate in panel A, the actual increase in effective sex ratio from 1950 to 1960 accounts for 23 percent (0.033 × 0.15 / [0.046 – 0.024]) of the average gender gap in mortality rates.[1]

[1] The change in imputed sex ratio in this period is about 0.15.

Figure 3.1 Five-year mortality rate for age cohort 50 to 54 by birth year and gender in Taiwan



Source: Authors' calculation.
Notes: The solid and dashed lines represent the five-year mortality rate for age 50 to 54, respectively, for men and women who were born between 1931 and 1950.

When the marriage ban was abolished, most veterans were older than. The majority of women of similar age had already gotten married. Thus, the veterans had to look for younger women as their wives. Consequently, the marriage age gap widened from about zero in the early 1950s to more than five years by 1970 (Chang and Zhang 2015). Considering the enlarging marriage age gap between husbands and wives, the sex ratio used in Table 3.1 may not accurately capture the actual matching of older husbands and younger wives. As a robustness check, in Table 3.2, we repeat the regressions in Table 3.1 using several alternative sex ratios. First, we define the sex ratio variable as the ratio of men aged 20 to 49 to women five years younger than them (15–44 years old). As shown in panel A, the findings in Table 3.1 largely hold. A higher sex ratio is associated with a higher male mortality rate. By contrast, there is no such relationship for women.

There is a concern about the wide age range in the definition of sex ratio. Even using 44 as a cutoff age for women in the above sex ratio variable, the sex ratio variable is still likely subject to measurement error. As a further robustness check, we narrow the age intervals used for calculating the sex ratio variable by 10 years. Specifically, we use the ratio of men aged 20 to 39 to women aged 20 to 39 as a measure of mating competition. The results in the second row are largely consistent with those in the first row. In the third set of regressions, we define the sex ratio variable as the number of men aged 20 to 39 relative to the number of women aged 15 to 34. In other words, women are five years younger than men in the calculation. As shown in the third row, the results closely mirror those in the second row. Men still suffer higher mortality.

The sex ratio variables used in Tables 3.1 and 3.2 are from Chang and Zhang (2015). Despite the authors' best effort, one may be concerned about whether the imputed effective sex ratio embodies any measurement error. As a robustness check, we opt not to use sex ratio variables. Instead we compare men and women born before and after 1939. Those born in 1940 were 19 years old in 1959, and most of them were not married by then; therefore, they were more affected by the policy change than were their older counterparts. We create a dummy variable for the cohort exposed to the change in marriage policy with a value of 1 if born after 1939 and 0 otherwise.

Table 3.1 Regressions of five-year mortality rate at age 50 to 54 on sex ratio at age 20

	(1) Men (mean = 0.046)	(2) Women (mean = 0.024)	(3) Difference (men – women)
	Panel A: Current residence		
Sex ratio	0.046***	0.013	0.033**
	(0.015)	(0.008)	(0.012)
n	420	420	
	Panel B: Residence five years ago		
Sex ratio	0.031**	0.008	0.023**
	(0.011)	(0.006)	(0.010)
n	420	420	

Source: Mortality rates are authors' calculations based on death registry records acquired from the Department of Health and the 1980 Population and Housing Census from the Directorate-General of Budget, Accounting and Statistics in Taiwan.

Notes: All estimates are derived from separate regressions. The sample consists of birth cohorts 1931 to 1950, who were 20 during 1951 to 1970. Dependent variable is county-level sex-cohort-specific five-year mortality rate at age 50 to 54, which is defined by the total deaths at age 50 to 54 for each cohort divided by the cohort size at the end of age 49. Sex ratio is the ratio of men to women of age 15 to 49 at the county level when the cohort reached age 20. In panel A, we use the current residence in the 1980 population census as the proxy for residence at age 20. In panel B, we use the residence in 1975 as the proxy for residence at age 20. All regressions include a full set of cohort dummies and county dummies. Robust standard errors clustered at the county level are in parentheses. **Significant at 5 percent level. ***Significant at 1 percent level.

Table 3.2 Regressions of five-year mortality rate at age 50 to 54 on alternative sex ratio

Sex ratio	(1) M (mean = 0.046)	(2) W (mean = 0.024)	(3) Difference (M – W)
M: 20–49/ W: 15–44	0.032*	0.008	0.0239
	(0.018)	(0.007)	(0.014)
M: 20–39/ W: 20–39	0.048***	0.018***	0.029***
	(0.009)	(0.005)	(0.007)
M: 20–39/ W: 15–34	0.046***	0.019***	0.028***
	(0.010)	(0.006)	(0.008)
n	420	420	

Source: Mortality rates are authors' own calculations based on death registry records acquired from the Department of Health and the 1980 Population and Housing Census from the Directorate-General of Budget, Accounting and Statistics in Taiwan.

Notes: All estimates are derived from separate regressions. The sample consists of birth cohorts 1931 to 1950, who were 20 during 1951 to 1970. Dependent variable is county-level sex-cohort-specific five-year mortality rate at age 50 to 54, which is defined by the total deaths at age 50 to 54 for each cohort divided by the cohort size at the end of age 49. Sex ratio is the ratio of men to women of indicated ages at the county level when the cohort reached age 20. We use the current residence in the 1980 population census as the proxy for residence at age 20. All regressions include a full set of cohort dummies and county dummies. Robust standard errors clustered at the county level are in parentheses. M = men; W = women. *Significant at 10 percent level. **Significant at 5 percent level. ***Significant at 1 percent level.

The regression in Table 3.3 includes a dummy variable for the cohort exposed to the marriage ban removal, a dummy for males, and their interaction term. The interaction term, representing the difference-in-differences estimate, is the variable of interest. The coefficient for the interaction term is 0.005, highly significant at the 1 percent level. Given the average mortality rates for men and women aged 50 to 54 is 0.046 and 0.024, exposure to the policy change widens the gender gap in mortality by 23 percent (0.005 / [0.046 – 0.024]). The impact is economically sizable. To a large extent, the effective sex ratio variable is exogenous. However, there is still a possibility that the variable is temporally correlated with some other factors, such as infectious disease and labor market situations, which matter to the outcome variable but are not controlled for in the regressions. Under the presence of a common trend between the sex ratio variable and the outcome variables, the sex ratio variable actually represents a trend inherent in the dependent variables across counties. In this case, if we randomly match the sex ratio variable with county codes and repeat Table 3.2, we still should be able to observe a positive coefficient for the sex ratio variable, which captures the common trend. Specifically, for each county, we first replace the sex ratio variable with the value of sex ratio variable from another randomly selected county and run regressions in Table 3.4. The process repeats 250 times. Table 3.4 presents the regression results of the falsification test based on the randomly shuffled sex ratio variable. None of the coefficient for the sex ratio variable in regressions on men and women is significant, largely dismissing the concern about an omitted common trend.

Table 3.3 Difference-in-differences regression of five-year mortality rate at age 50 to 54

Male	0.017***
	(0.001)
Dummy for cohort 1940–1950	–0.008***
	(0.001)
Male × Dummy for Cohort 1940–1950	0.005***
	(0.001)
n	840

Source: Mortality rates are authors' own calculations based on death registry records acquired from the Department of Health and the 1980 Population and Housing Census from the Directorate-General of Budget, Accounting and Statistics in Taiwan.

Notes: The sample covers birth cohort 1931 to 1950, who were 20 during 1951 to 1970. Dependent variable is county-level sex-cohort-specific five-year mortality rate at age 50-54, which is defined by the total deaths at age 50 to 54 for each cohort divided by the cohort size at the end of age 49. In addition to the gender and cohort dummy and their interaction term, the regression includes a full set of county dummies. Robust standard errors clustered at the county level are in parentheses. ***Significant at 1 percent level.

Table 3.4 Regressions of five-year mortality rate at age 50 to 54 on randomly assigned sex ratio at age 20

	(1) Men (mean = 0.046)	(2) Women (mean = 0.024)	(3) Difference (men – women)
	Panel A: Current residence		
Randomly assigned sex ratio	–0.0002060	0.0000846	–0.0002908
	(0.0035640)	(0.0020280)	(0.0041091)
	Panel B: Residence five years ago		
Randomly assigned sex ratio	0.0000992	0.0000159	0.0000834
	(0.0029076)	(0.0017863)	(0.0034910)

Source: Mortality rates are authors' own calculations based on death registry records acquired from the Department of Health and the 1980 Population and Housing Census from the Directorate-General of Budget, Accounting and Statistics in Taiwan.

Notes: The sample covers birth cohort 1931 to 1950, who were 20 during 1951 to 1970. Dependent variable is county-level sex-cohort-specific five-year mortality rate at age 50 to 54, which is defined by the total deaths at age 50 to 54 for each cohort divided by the cohort size at the end of age 49. Sex ratio is the ratio of men to women of age 15 to 49 at the county level when the cohort reached age 20. In each regression, county sex ratios at age 20 were randomly shuffled for each birth cohort. The process was repeated 250 times to acquire 250 estimates for each sex/cohort. The means of the estimates are reported. The standard deviations of the estimates are in parentheses.

Despite the various robustness checks, our estimates of the sex ratio effect on mortality still face two issues. First, due to a lack of information, we are not able to separate the veterans from Mainland China from local Taiwanese people in the mortality data. It is thus likely that the average veteran from Mainland China would have had experienced a higher mortality rate than the average local Taiwanese man simply because the former's war experience. Meanwhile, the veterans were also the main driver of the elevated sex ratio. Second, an increase in the sex ratio also means more "losers" in the marriage competition, who remained unmarried in the end. The outcome of the marriage competition itself can lead to inequality in mortality between married and unmarried men because married men are known to have lower mortality than men of other marital status. In the following analysis, we utilize individual level data, which allow us to control for marital status and birth place information to mitigate these two concerns.

Impact on Self-Reported Health Status and Depression Based on Elderly Survey Data

In addition to the aforementioned issues, the analyses based on aggregate mortality data at the county level have a drawback in that individual behavioral factors cannot be controlled for. Considering that individual behavior, such as smoking, is a major risk factor for morbidity and mortality (Case and Paxson 2005), we further look at the impact of mating competition on the gender difference in self-reported health status and depression measured at the individual level using the Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan in 1996 and 2003. We estimate the following linear model.

$$Y_i^g = S_i\beta^g + X_i\theta^g + \varepsilon_i^g, \qquad (2)$$

where Y_i^g denotes an indicator variable of self-reported health status or an index of depression for individual i; g represents gender (male and female).

Apart from the sex ratio variable S_i, the regressions also control for X_i, which is a vector consisting of age, age squared, ethnicity, education, marital status, lifestyle variables such as smoking, drinking alcohol, chewing betel nuts, and frequency of exercise, a full set of birth year fixed effects, and a dummy indicating sample 2003.

Columns (1) and (2) in Table 3.5 present the regression results of two health outcome variables (self-reported health status and depression measure) on sex ratio at age 20 for elderly men and women, respectively. Columns (3) and (6) report the difference in the coefficient for the sex ratio variable between the first two regressions for each outcome variable. Columns (1) through (3) report estimates of self-reported good health. Panel A is for the full sample. Sex ratios that women faced at age 20 are highly correlated with their reported good health status in later life, while the impact for men is not statistically different from zero. The gender difference in self-reported health status is significant at the 10 percent level. Exposure to a higher sex ratio during young adulthood significantly enlarges the gender gap in self-reported health status. The Center for Epidemiological Studies Depression measure is a sum of 10 depression items, each of which scores from 0 to 3, with a higher value indicating higher frequency of depression. Columns (4) through (6) examine the impact of sex ratio on the Center for Epidemiological Studies Depression measure. Different sex ratios at age 20 have drastically different impacts on men's and women's depression measures. Men's depression score is positively associated with the effective sex ratio for cohort 15 to 49 when that cohort entered the marriage market at age 20. Conversely, a higher sex ratio at age 20 greatly lowers women's depression. An increase in sex ratio by 10 percentage points would widen the gender difference in depression scores by 0.85. Given the mean gender difference is only 0.84, the increase in mating competition essentially explains the entire gender gap.

Panel A includes both single and married people. Although in panel A we have controlled for marital status, the regression results hinge on a key assumption that all other variables affect married and single people's health outcomes in the same way. Since marital status is highly positively associated with health status, the above findings may be driven by the existence of excess single men inherent in the definition of an imbalanced sex ratio (a higher value means more men relative to women). To address this concern, in panel B, we restrict the sample to married people only and repeat the exercises in panel A. In doing so, we avoid the problem of sample selection. In a region with a higher ratio of men to women of marriageable age, mathematically more men would end up being single. A comparison of panel B and panel A indicates that the main findings do not change much. Exposure to higher sex ratios at age 20 improves women's likelihood of good health status (column [2]) and reduces their chance of depression (column [5]), while that same exposure is detrimental to men's mental health.

Table 3.5 Regressions of health outcomes of the elderly on sex ratio at age 20

	Good/very good health			***Depression***		
	(1) Men	**(2) Women**	**(3) Difference**	**(4) Men**	**(5) Women**	**(6) Difference**
	Panel A: Full sample					
	(mean = 0.510)	(mean = 0.419)		(mean = 4.070)	(mean = 4.915)	
Sex ratio	0.043	0.309***	–0.266*	3.752*	–4.692**	8.445***
	(0.136)	(0.088)	(0.154)	(1.867)	(1.763)	(2.194)
n	2,081	1,980		1,990	1,899	
	Panel B: Only married elderly individuals					
	(mean = 0.524)	(mean = 0.431)		(mean = 3.751)	(mean = 4.681)	
Sex ratio	0.078	0.303***	–0.224	3.401*	–5.745**	9.146***
	(0.167)	(0.105)	(0.187)	(1.851)	(2.415)	(1.927)
n	1,808	1,570		1,736	1,510	
	Panel C: Only married elderly individuals born in Taiwan					
	(mean = 0.530)	(mean = 0.431)		(mean = 3.750)	(mean = 4.667)	
Sex ratio	0.058	0.273**	–0.215	3.512	–5.519	9.031***
	(0.173)	(0.112)	(0.185)	(2.066)	(2.615)	(2.265)
n	1,685	1,529		1,619	1,476	

Source: Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan, 1996 and 2003 waves.

Notes: We use two samples drawn in 1996 and 2003, respectively. The 1996 sample consists of elderly born between 1926 and 1946. The 2003 sample contains elderly born between 1946 and 1953. We combine the two samples. Note that their health outcomes are measured in 1996 and 2003, respectively. The dependent variable in columns (1) to (2) is a dummy variable indicating whether an elderly individual self-reported very good or good health based on a five-point scale including very good, good, fair, poor, and very poor. The dependent variable in columns (4) to (5) is the depression measure (Center for Epidemiological Studies Depression), which is a sum of 10 depression items, each of which scores from 0 to 3, with a higher value indicating higher frequency of depression. The possible maximum score is thus 30. Columns (3) and (6) report the difference between men and women. Sex ratio is the ratio of men to women of age 15 to 49 in the county when one reached age 20. Married individuals are individuals who are married and whose spouses are still alive. All estimates are from separate regressions. All regressions control for age; age squared; ethnicity; education; marital status (panel A); health behavior variables such as drinking, smoking, exercising, and chewing betel nuts; a dummy indicating the sample from the 2003 wave; and a full set of county dummies. Robust standard errors clustered at the county level are in parentheses. *Significant at 10 percent level. **Significant at 5 percent level. ***Significant at 1 percent level.

One may be concerned about inherent differences between native Taiwanese people and immigrants from Mainland China, war-torn veterans in particular. In panel C of Table 3.5, we drop individuals born in Mainland China. The main results hold. We further check the robustness of the findings in Table 3.5 by using alternative sex ratios similar to those in Table 3.2 and report the regression results for the sex ratio variables in Table 3.6. When using alternative sex ratios at age 20, the coefficient for the gender difference in self-reported good health loses significance, while exposure to skewed sex ratios yields a significant gender difference in depression scores. The results are robust to regressions based on the full sample, a subsample of married people, or a subsample of married people born in Taiwan.

Table 3.6 Regressions of health outcomes of the elderly on alternative sex ratios

Sex ratios	*Good/very good health* (1) Men	(2) Women	(3) Difference	*Depression* (4) Men	(5) Women	(6) Difference
			Panel A: Full sample			
M: 20–49/ W: 15–44	0.044	0.300***	–0.256	4.076**	–4.417**	8.493***
	(0.132)	(0.089)	(0.149)	(1.702)	(1.646)	(2.004)
M: 20–39/ W: 20–39	0.036	0.174*	–0.138	1.670	–2.871*	4.541**
	(0.110)	(0.090)	(0.119)	(1.309)	(1.397)	(1.620)
M: 20–39/ W: 15–34	0.072	0.203**	–0.131	2.154	–2.920*	5.074***
	(0.114)	(0.094)	(0.139)	(1.312)	(1.528)	(1.772)
n	2,081	1,980		1,990	1,899	
			Panel B: Only married elderly individuals			
M: 20–49/ W: 15–44	0.081	0.287***	–0.207	3.816**	–5.493**	9.309***
	(0.162)	(0.097)	(0.178)	(1.779)	(2.196)	(1.940)
M: 20–39/ W: 20–39	0.058	0.174	–0.115	1.677	–3.493**	5.170***
	(0.120)	(0.105)	(0.135)	(1.354)	(1.645)	(1.451)
M: 20–39/ W: 15–34	0.095	0.212*	–0.117	2.350	–3.686**	6.036***
	(0.130)	(0.106)	(0.159)	(1.373)	(1.761)	(1.723)
n	1,808	1,570		1,736	1,510	
			Panel C: Only married elderly individuals born in Taiwan			
M: 20–49/ W: 15–44	0.078	0.263**	–0.185	3.998*	–5.351**	9.349***
	(0.161)	(0.108)	(0.173)	(2.006)	(2.377)	(2.272)
M: 20–39/ W: 20–39	0.045	0.150	–0.106	1.587	–3.307*	4.894***
	(0.130)	(0.110)	(0.132)	(1.530)	(1.814)	(1.698)
M: 20–39/ W: 15–34	0.095	0.191	–0.096	2.301	–3.541*	5.843***
	(0.133)	(0.112)	(0.155)	(1.584)	(1.925)	(1.998)
n	1,685	1,529		1,619	1,476	

Source: Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan, 1996 and 2003 waves.

Notes: We use two samples drawn in 1996 and 2003, respectively. The 1996 sample consists of elderly born between 1926 and 1946. The 2003 sample contains elderly born between 1946 and 1953. We combine the two samples. Note that their health outcomes are measured in 1996 and 2003, respectively. All regression specifications are the same as in Table 3.5 except for the county sex ratios. Robust standard errors clustered at the county level are in parentheses. M = men; W = women. *Significant at 10 percent level. **Significant at 5 percent level. ***Significant at 1 percent level.

Table 3.7 reports the heterogeneous effects of mating competition on health outcomes by education level. Three education dummy variables, no schooling, receiving basic education (1–12 years of schooling) or higher education (more than 12 years of schooling), are created. The reference group is individuals with no schooling. In addition to the sex ratio variable and other control variables used in Table 3.5, the two education variables ("1–12 years of schooling" and "more than 12 years of schooling") and their interactions with the sex ratio variable are included in regressions. Regardless of education level, a higher sex ratio increases women's self-reported good health status. The effect becomes stronger as women's education level rises. For illiterate men, local sex ratio alone does not seem to matter much to self-reported health status as shown by the insignificant coefficient for the sex ratio variable. However, facing the same sex ratio, educated men report better health status than do less educated men.

Sex ratio shapes men's and women's scores on the Center for Epidemiological Studies Depression measure differently. A high sex ratio increases men's likelihood of depression across education levels while lowering the level of women's depressive symptoms. Interesting to note, the impact of sex ratio on reducing the likelihood of depression increases with women's education level but not with men's.

Table 3.7 Heterogeneous sex ratio effects on health outcomes, by education

	Good/very good health		*Depression*	
	(1) Men	**(2) Women**	**(3) Men**	**(4) Women**
Sex ratio	–0.051	0.230**	4.064**	–4.080**
	(0.155)	(0.080)	(1.841)	(1.822)
Sex Ratio × (Education = 1–12 Years)	0.079*	0.098***	–0.281	–0.741**
	(0.043)	(0.014)	(0.492)	(0.322)
Sex Ratio × (Education ≥ 13 years)	0.204**	0.148***	–0.877	–1.939***
	(0.058)	(0.047)	(0.518)	(0.485)
n	2,081	1,980	1,990	1,899

Source: Survey of Health and Living Status of the Middle-aged and the Elderly in Taiwan, 1996 and 2003 waves.

Notes: We use two samples drawn in 1996 and 2003, respectively. The 1996 sample consists of elderly born between 1926 and 1946. The 2003 sample contains elderly born between 1946 and 1953. We combine the two samples. Note that their health outcomes are measured in 1996 and 2003, respectively. The reference group is individuals with no schooling. All regressions control for age; age squared; ethnicity; marital status; health behavior variables such as drinking, smoking, exercising, and chewing betel nuts; a dummy indicating the sample from the 2003 wave; and a full set of county dummies. Robust standard errors clustered at the county level are in parentheses. *Significant at 10 percent level. **Significant at 5 percent level. ***Significant at 1 percent level.

Possible Channels

Our main story is that when there were more men relative to women in the marriage market, women became choosier. Holding everything else constant, women prefer to marry wealthier men (Buss 1989; Chiappori, Oreffice and Quintana-Domeque 2012). Knowing women's preference for wealth, young men or their parents respond by using various means to accumulate more wealth. Engaging in risky jobs, such as being an entrepreneur, and working longer hours are two obvious options to get rich. Chang and Zhang (2015) establish the link that mating competition spurs men's entrepreneurship using the same natural experiment in Taiwan. Here we provide more evidence about the second option. Men exposed to elevated marriage market competition work longer hours than women and are therefore more likely to feel stress.

Table 3.8 presents the separate regressions of work status and work hours in the past week for men and women on the effective sex ratios facing them at age 20. As shown in panel A, higher sex ratios prompt men to work and reduce women's pressure to work. A 10 percentage point increase in sex ratio would increase men's labor force participation rate by 3 percentage points and decrease women's labor force participation rate by 3.6 percentage points. Given that only 43.5 percent of women worked, the effect is sizable. Such an increase in sex ratio accounts for 18 percent (0.653 × 0.1 / [0.797 – 0.435]) of the observed gender gap in labor force participation rate. The estimates of work hours in the past week are reported in panel B. As in panel A, higher sex ratios induce men to work longer hours, while women can afford to work shorter hours. If sex ratios increased by 10 percentage points, men would work 1.7 extra hours, while women would work 1.4 hours less. A 10 percentage point increase in sex ratio explains 15.3 percent of the average gender gap in work hours.

Table 3.8 Regressions of labor market outcomes on sex ratio at age 20

	Panel A: Worked last week		
	(1) Men (mean = 0.797)	(2) Women (mean = 0.435)	(3) Difference (men – women)
Sex ratio	0.290***	–0.363***	0.653***
	(0.093)	(0.090)	(0.167)
n	181,224	164,195	
	Panel B: Work hours last week		
	(1) Men (mean = 40.353)	**(2) Women (mean = 19.707)**	**(3) Difference (men – women)**
Sex ratio	17.336***	–14.252***	31.588***
	(5.197)	(4.263)	(8.778)
n	181,224	164,195	

Source: Taiwan Manpower Survey 1978 (DGBAS 2014).

Notes: Dependent variable in panel A is a dummy variable indicating whether one worked in the week prior to the survey. Dependent variable in panel B is the total work hours in the week prior to the survey. Sex ratio is the ratio of men to women of age 15 to 49 in the county when the respondent reached age 20. The sample consists of individuals born between 1928 and 1958 (age 20–50 in 1978). All regressions control for age, age squared, marital status, education, household size, county-level share of prime age (20–64) population at age 20, log of prime age male population at age 20, and a full set of county dummies. Robust standard errors clustered at the county level are in parentheses. ***Significant at 1 percent level.

As men put more effort into work, on the one hand, their wealth may increase, and so does their utility; on the other hand their leisure time is squeezed, making them less happy. So estimating the net effect of a greater sex ratio on utility is an empirical question. Table 3.9 presents some evidence of the impact of sex ratio imbalances on life satisfaction, a common measure of utility. As indicated in column (1), a higher sex ratio is associated with significantly lower life satisfaction for men. By comparison, women's happiness increases as marriage market conditions become more favorable to women. A 10 percent increase in sex ratio accounts for almost all the observed gender disparity in life satisfaction (97 percent = 0.029 / 0.03). It is likely that sex ratio imbalance exerts some negative consequences on men's health outcomes through the channel of lowered life satisfaction.

Table 3.9 Regression of life satisfaction on sex ratio at age 20

	(1) Men (mean = 0.80)	(2) Women (mean = 0.83)	(3) Difference (men – women)
Sex ratio	–0.16*	0.13	–0.29*
	(0.09)	(0.13)	(0.16)
Married	0.04	0.07**	–0.04
	(0.03)	(0.03)	(0.04)
n	3,445	2,311	

Source: National Survey of Living Conditions in 1992, 1993, and 1994 (MoI various years).

Notes: All three waves, 1992, 1993, and 1994, are pooled. We restrict the sample to cohorts 1926 to 1960. Dependent variable is a dummy variable indicating that one is satisfied, very satisfied, or extremely satisfied about his or her current life based on a six-point scale from *extremely satisfied* to *extremely unsatisfied.* Sex ratio is the ratio of men to women of age 15 to 49 in the county when the respondent reached age 20. All regressions control for age, age squared, education, a full set of income dummies, county dummies, and two dummies indicating 1993 and 1994 waves. Robust standard errors clustered at the county level are in parentheses. *Significant at 10 percent level. **Significant at 5 percent level.

4. CONCLUSIONS

This paper complements the existing literature on gender difference in competition by focusing on the negative long-term health impact, including mortality. Using the removal of the marriage ban on soldiers from Mainland China in Taiwan as a natural experiment, we show that their exposure to the sudden elevation in mating competition in early adulthood exerted long-term negative health consequences on men. Facing greater competition, men ended up with a higher likelihood of depression and mortality. Women, on the contrary, enjoyed greater bargaining power (and thus better health outcomes) because they had abundant men from whom to choose.

It has been well documented that men are generally more competitive than women. As shown in our paper, the gender difference in competiveness may contribute to the gender gap in mortality. Exposure to intense competitive pressures can come with heavy health and mortality costs. There is a hidden health cost associated with competition.

High sex ratios have persisted in some Asian countries, including China and India, for a few decades. As the excess men age in the next several decades, the toll on their health as a result of overcompetition in the marriage market will likely exhibit.

APPENDIX: CALCULATION OF MORTALITY RATE

The cohort-level mortality rate M_{gcbt} is computed using the 1980 population census data and the 1981 to 2004 death registry as follows.

$$M_{gcbt} = \frac{D_{gcbt}}{P_{gcbt}},$$

where D_{gcbt} is the number of individuals belonging to cohort gcb who died in period t (which indicates a five-year interval), with g indexing gender, c indexing county of residence, and b indexing birth cohort, and P_{gcbt} is the number of individuals belonging to cohort gcb at the beginning of period t. Number of deaths D_{gcbt} is computed using the death registry. Population size P_{gcbt} is computed using the 1980 population census, which, enumerated at the end of 1980, gives the number of individuals at the beginning of 1981. Using the number of individuals belonging to a cohort who died (taken from the death registry) in years prior to 1980, we impute the population size of that cohort in years prior to 1980.

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