Gender Imbalance, Marriage Stability, and Divorce Rate: Evidence from China*

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Abstract

The deficit of men or women in a regional marriage market is a commonly observed phenomenon stemming from factors such as “Missing Girls” at birth, immigration, and higher mortality rates among men due to war. However, the impact of this deficit on marriage stability remains not well understood. In this paper, using provincial, census, and household survey data in China, we find that a higher male-to-female ratio increases divorce rates. Further analyses support the hypothesis that this impact is primarily driven by married women having more options outside their marriage. The effect is more pronounced in economies with greater income inequality, where there are more wealthy prospective partners. These findings highlight the significance of gender balance in sustaining stable marriages and uncover a new contributing factor to the escalating divorce rates in China.

Keywords: Divorce Rate, Gender Imbalance, Outside Option, Inequality
JEL Codes: D10, J12, J16, N35

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1 Introduction

The phenomenon of "Missing Women" caused by son preference is a stark issue in many Asian countries (Sen, 1990, 1992).[1] The global count of "Missing Women" is expected to peak at 150 million in 2035 (Bongaarts and Guilmoto, 2015). The deficit of women has been documented to affect various aspects of social development profoundly. However, its impact on marriage stability has been largely ignored in the literature. Addressing this aspect could provide a deeper understanding of the broader societal impacts of gender imbalance.

The importance of a stable marriage to the welfare of household members and the overall societal development cannot be overstated. Regrettably, many developing economies witnessed a notable increase in divorce rates over the past decades.[2] In particular, the crude divorce rates[3] in China rose from 0.96 in 2000 to 3.36 in 2019, representing a nearly 3.5-fold surge.[4] In fact, the crude divorce rate in China has surpassed that of the United States since 2016, prompting significant government attention and the implementation of various policies. It is noteworthy that the surge in divorce rates has coincided with the marriage and divorce of the birth-control generation, which has faced a highly imbalanced sex ratio. This paper investigates whether the sex imbalance has contributed to the escalating divorce rate in China.

The impact of sex ratio on divorce is ex-ante ambiguous. Previous literature has provided evidence that women can benefit in a marriage market characterized by a relative shortage of females. For example, a higher male-to-female ratio can boost women's chances of marrying up, raise bride prices, and increase women's within-household bargaining power. Theoretically, this would enhance their satisfaction in marriage and increase marital stability. On the other hand, gender imbalance in the marriage market can create more "outside options" for married women. In other words, there are more available men in the marriage market, which may prompt married women to consider leaving their current marriages. We present a simple model in Appendix Section A to demonstrate the impacts of the two forces, showing that it is possible to derive either positive or negative overall impacts of sex imbalance on divorce rates from the same theoretical framework under varying conditions.

[1] The "missing women" phenomenon refers to the lower number of women in the population than expected if both genders were subject to similar birth and mortality rates.

[2] For example, according to the OECD Family Database, the crude divorce rates rose from 0.66 to 1.8 in Vietnam and from 1.1 to 1.9 in Thailand from 2000 to 2019 (OECD and Centre, 2021).

[3] The crude divorce rate is the number of divorce cases per 1000 people during a given year.

[4] Before 2001, divorce was very rare in China as women's property rights were not well-protected after divorce. In 2001, the government launched a pro-women divorce reform, which liberalized divorce in favor of women and secured women's property rights after separation (Sun and Zhao, 2016).

There are multiple reasons why more "outside options" could potentially increase the likelihood of married women seeking a divorce. In the perfect information setting, there would be no divorce, as people who want a marriage would always marry their best option and stay in the marriage forever. However, at least two sources of imperfect information contribute to the possibility of divorce. Firstly, the information on which couples relied before entering into marriage is incomplete (Becker et al., 1977). Once married, if the wife finds the realized utility low, such as in cases of an abusive husband, she may feel inclined to explore alternative prospects outside the current marriage. Secondly, changes in the choice set, such as the rise of new contacts in the workplace, can lead to divorce (McKinnish, 2004). If a married woman meets a man she believes can provide a better marriage than her current husband, she may divorce and remarry the new prospect. In both scenarios, more available men in the marriage market expand the options for married women seeking to improve their marital situation.

The outside option channel is particularly relevant when there is sex imbalance in regional marriage markets—a phenomenon frequently observed. Aside from "missing women," various factors can lead to sex imbalance. For example, the regional industrial structure can shape the gender composition of laborers, which subsequently affects the gender structure in the marriage market (Kearney and Wilson, 2018). Additionally, immigration (Angrist, 2002; Lafortune, 2013) and war (Abramitzky et al., 2011; Brainerd, 2017; Ogasawara and Igarashi, 2021) have been documented in previous literature as factors that impact sex imbalance. Moreover, as more women migrate to urban areas in search of high-income husbands, simultaneous imbalances can arise, with a deficit of men in the urban marriage market and a deficit of women in the rural marriage market (Edlund, 2005; Ong et al., 2020). By documenting the outside option channel, our framework offers a valuable perspective for understanding the impact of sex imbalance on marriage stability.

Despite its perceived importance, the outside option channel has received very little empirical attention, possibly due to data limitations. Our data uniquely provides information on remarriage and the quality of the subsequent husbands, offering an opportunity to fill this gap in the literature by investigating its implications.

We examine the hypotheses in the Chinese context. China suffers a bigger challenge than other countries from sex imbalance because it faces a severe "Missing Women" problem and because of its huge population size.[5] China has one of the highest sex ratios at birth in the world, with 121.7 boys born for every 100 girls in rural areas in 2000 (Li, 2007). The nationwide sex imbalance is believed to be predominantly driven by the

[5]See Das Gupta et al. (2010) for a discussion of the challenges confronted by China because of the gender imbalance.

combination of son preference and birth control policies. The deeply rooted preference for sons stems from their expected role in caring for parents in old age and continuing the family lineage. Consequently, when China implemented the One-Child Policy (OCP) in the late 1970s, there was a significant surge in parents' tendency to practice sex selection.[6] The widespread utilization of B-ultrasound devices greatly facilitated the sex selection for many households.[7] OCP remained in effect until a relaxation in 2011 and was formally abandoned in 2016.[8] The policy has had a lasting impact on cohorts spanning over 30 years and is expected to continue influencing the marriage market for decades.

While commonly referred to as the "one-child policy," it's important to note that additional births were not always prohibited, and the stringency of enforcement varied substantially across regions and over time. As a prominent fertility policy, the fine rates for above-quota births, set by provincial governments based on their local conditions and changed over time, ranged from 10 to 50 percent wage deductions spanning 5 to 14 years (Ebenstein, 2010). Furthermore, Qian (2009) and Liu (2014) both point out that within the same province, there were substantial variations in local fertility policies. Even with the same policy, local implementation exhibited differences due to factors such as varying administrative capacities and implementation approaches (Li and Zhang, 2017). In this paper, we leverage the extensive geographical and temporal variation for identification. Specifically, for the main analyses using individual-level data, we construct the sex ratio at the prefecture times birth year level and examine its impact on divorce, controlling for various individual-level characteristics, birth year fixed effects, and prefecture fixed effects. We also conduct 2SLS estimations using several instrumental variables for robustness.

Our study constitutes three sets of analyses. Firstly, using province-level panel data, three waves of individual-level census data, and individual-level survey data, we

[6]Studies based on Chinese census data suggest that the "missing girls" phenomenon is closely linked to the enforcement of the OCP. For example, Ebenstein (2010) presents evidence from Chinese census data that the "missing girls" are caused by the enforcement of the One-Child Policy (OCP). Li et al. (2011) find that the OCP could explain 94%, 57%, and 54% of total increases in sex ratios for the 1980s, the 1990s, and the 2001-2005 birth cohorts, respectively. More detailed discussions about the OCP in China can be found in Bongaarts and Greenhalgh (1985) and Greenhalgh (2005), among others.

[7]Chen et al. (2013) estimate that roughly 40% to 50% of the increase in the sex imbalance at birth can be explained by local access to B-ultrasound examination. Similarly, according to Hesketh and Xing (2006)'s estimation, the number of missing women attributable to gender selection using B-ultrasound devices was up to 80 million in China and India.

[8]The first relaxation of the policy occurred in 2011 when it became permissible for couples who were both only children to have two children. On January 1, 2016, the revised Article 18, Clause 1 of the "People's Republic of China Population and Family Planning Law" stated that the state encourages couples to have two children, effectively ending the one-child policy that had been in place for more than three decades.

demonstrate that a surplus of males in the marriage market leads to an increase in the divorce rate. Specifically, at the province level, an increase in the sex ratio (male-to-female ratio) among populations aged 20-40 significantly raises the annual divorce rate. For individual-level analyses, we examine the census sample data in 2005, 2010, and 2015. We document that the sex ratio faced by young women at the prefecture times birth year level significantly increases their probability of being currently divorced. To complement the census data, we utilize marriage history information from the 2010 wave of the China Family Panel Studies (CFPS) data and construct the "ever divorced" indicator as the outcome variable. We find that a higher sex ratio increases the likelihood of women experiencing a divorce at some point in their lives. Moreover, we rule out several other potential explanations, such as changes in the age of marriage, increased cohabitation with the husband's parents, and shifts in the gender education gap. There were no significant changes in divorce laws that could potentially introduce confounding factors to our results during our sample period.[9]

The second set of analyses examines the underlying channel. We first demonstrate that, as documented in the existing literature, a higher sex ratio improves the quality of marriage for women. These findings indicate that the impact of sex ratio on divorce is not due to women facing difficulties in finding suitable partners for their first marriage. Specifically, we find that women exposed to a higher sex ratio are married to men with better education and higher income, and rural *Hukou* women are more likely to have husbands with urban *Hukou*. Additionally, these women live in higher-quality houses characterized by larger areas, more rooms, and higher house prices. Moreover, in terms of mental well-being, women facing a higher sex ratio report lower levels of depression.

Further evidence supports the importance of the outside option channel. We observe that divorced or widowed women in local marriage markets with a higher sex ratio have a higher probability of remarriage. Additionally, women who remarry in male-skewed markets tend to have higher-quality remarriage partners. A higher likelihood of remarriage and the expectation of a high-quality subsequent husband can reduce the reluctance to divorce, explaining the higher divorce rates observed in markets with an excess of men.

Furthermore, we provide evidence that the impact of sex ratio on divorce is intensified in a more unequal economy. Assuming wealth is crucial when women choose a marriage

[9]The pro-women divorce reform in 2001 was the major relevant change in the marriage law in recent decades. Notably, between 2005 and 2015, there were no considerable alterations in divorce laws, except for the introduction of the "Judicial Interpretation III of the Marriage Law" in 2011, which undermines the protection of women's property rights post-divorce. As our regressions control for calendar year fixed effects and birth year fixed effects, Interpretation III's simultaneous nationwide application should not affect our results.

partner, a married woman who encounters a wealthier man than her current husband may choose to divorce and remarry. With a constant number of potential mates, higher inequality means a larger fraction of men with high incomes or wealth, providing more options for married women. Therefore, we argue that income inequality can amplify the impact of sex ratio on divorce, as married women have more wealthy outside options.

The third set of analyses examines the impacts of sex imbalance on outcomes for men. We observe increased divorce rates when men face higher male-to-female ratios. Additionally, we find that men spend more time caring for family members during weekdays, likely due to increased pressure to contribute more at home. Evidence on men's mental health suggests they suffer more in these situations. Overall, the results indicate that in a high sex ratio environment, men's marital surplus is diminished.

This paper contributes to several strands of literature. First, this paper mainly contributes to the large strand of literature that examines the impacts of gender imbalance since Becker's seminal work (Becker, 1973, 1981), especially the consequence of "Missing Girls" (Sen, 1990, 1992). This literature has overlooked the potential destabilizing effects of a deficit of marriageable women on marriages, and this paper aims to fill this gap. Existing studies have focused on the impacts of sex imbalance on marriage decisions, within-household resource allocation, labor market outcomes, and social stability.[10] Specifically, it has been documented that the relative shortage of marriageable women increases their probability of marrying up (Das Gupta et al., 2010; Abramitzky et al., 2011; Du et al., 2015), boosts the bride price (Francis, 2011; Tertilt, 2005; Nick and Walsh, 2007; Grossbard, 2015), and pushes grooms or grooms' parents to save more and purchase high-value houses/apartments for their marriage (Wei and Zhang, 2011a; Wei et al., 2017; Nie, 2020; Tan et al., 2021; Bhaskar et al., 2023). As a natural extension, sex imbalance could significantly influence marriage stability. This paper examines the impact and investigates the underlying channel.

Few studies have explored the relationship between the deficit of women and divorce rates. Most of the related papers examine the symmetric case where there are "missing men" in the marriage market due to high male mortality rates during wars. Existing papers primarily concentrate on developed countries. For example, Abramitzky et al. (2011) document that the relative shortage of men after WWI in France reduced their divorce rates. On the contrary, Brainerd (2017) and Ogasawara and Igarashi (2021)

[10] For papers on the impacts of sex imbalance on marriage and labor market outcomes, see, for example, Grossbard-Shechtman (1984), Samuelson (1985), Angrist (2002), Chiappori et al. (2002), Lafortune (2013), Negrusa and Negrusa (2014), Ong et al. (2020). For impacts on crimes such as sexual assault, violent crimes, and trafficking in females, see, for example, Hudson and Den Boer (2002, 2004), Hesketh and Xing (2006), Edlund et al. (2013), Cameron et al. (2019). Gender imbalance also changes parents' investment in children's education, but the results are mixed in the literature (Bhaskar and Hopkins, 2016; Lafortune, 2013; Bhaskar et al., 2023).

provide evidence from Russia and Japan, respectively, that male scarcity due to WWII increased divorce rates. Using war to identify the causal effect of gender imbalance on divorce rates faces potential challenges. War may cause mental health problems, which could influence the divorce decisions of veterans.[11] War can also alter labor market conditions for both men and women, which can influence marriage outcomes. Our paper examines a different setting with the opposite scenario where there are more men than women. Given the differences in what men and women seek from marriage and the varying contexts between developed and developing countries, this paper contributes to a deeper understanding of the relationship between gender imbalance and marriage outcomes. The context of China, known for a stark increase in divorce rate in recent years and a strong son preference, makes it a useful step towards explaining the complex interplay between gender norms and marriage market outcomes.

Additionally, several early sociology studies document the correlation between the sex ratio in the marriage market and divorce rate (Trent and South, 1989; South and Lloyd, 1992, 1995).[12] We complement by exploiting more comprehensive data with larger sample sizes, which enable us to examine causality and explore the underlying mechanisms.

Second, this paper contributes to the literature investigating determinants of divorce. A rich literature from demographers, sociologists and economists documents the effects of various factors such as education (e.g., Bruze et al., 2015), age of marriage (Teachman, 2002; Rotz, 2016), employment, income (see, for example, Bertrand et al., 2015), presence of a child and the gender of the child (e.g., Dahl and Moretti, 2008; Mammen, 2008), culture (Hirschman and Teerawichitchainan, 2003; Furtado et al., 2013), etc.[13]

[11]Particularly, Negrusa et al. (2014) find that deployment in the Iraq and Afghanistan wars significantly increases American soldiers' divorce rate. For more papers on the impacts of war on mental health, see, e.g., Angrist and Johnson IV (2000); Negrusa and Negrusa (2014).

[12]Additionally, two papers address the impacts of oversupply on divorce on one side of the marriage market, although their focus differs from ours. Lafortune (2013) shows that an exogenous rising supply of males relative to females due to immigration destabilizes males' marriages. Their research centers on the impacts of sex imbalance on pre-marriage investment changes. Weiss et al. (2013) investigate the handover of Hong Kong to the People's Republic of China in 1997, which they take as a supply shock of low-education mainland women. They find that more Hong Kong men have turned to marrying mainland women, while Hong Kong women have lower marriage rates, higher divorce rates, and increased emigration. Their focus is not the sex imbalance but the change in the composition of marriageable women.

[13]Other factors include duration of the marriage (Gottman and Levenson, 2000), communication difficulties between wife and husband (Thompson, 2008), marriage and divorce law (Kneip et al., 2014; Brassiolo, 2016), race (Teachman, 2002; Zhang and Van Hook, 2009), religion (Teachman, 2002; Vaaler et al., 2009; Glass and Levchak, 2014), public policy (Barham et al., 2006; Tjøtta and Vaage, 2008), discrimination in the labor market (Teachman and Tedrow, 2008), experience of parent's divorce/intergeneration transmission of divorce (Teachman, 2002; Lyngstad and Engelhardt, 2009; Fu and Wolfinger, 2011; Wolfinger, 2011; Sieben and Verbakel, 2013), and risk preference (Light and Ahn, 2010). The literature review on the linkage between social, demographic, and economic variables and divorce can be found in Faust and McKibben (1999), White (1990), and Amato (2010).

However, the impact of gender imbalance in the marriage market on divorce rates or marriage stability is ignored to a large extent by economic literature.[14] This paper highlights that the increasing divorce rates in China can be partially attributed to the deficit of marriageable women in the marriage market. Empirical evidence is consistent with the notion that the main driving force is the increased outside options available to married women, which may also be relevant in other economies. This paper calls for further investigation into the potential impacts of gender imbalance in regional marriage markets shaped by factors such as industrial structure, labor migration, or urbanization observed in other settings.

Third, this paper contributes to the literature investigating the impacts of inequality on social stability. For example, previous literature shows that inequality affects crimes (e.g., Kang, 2016) and sociopolitical unrest (Venieris and Gupta, 1986; Alesina and Perotti, 1996; Perotti, 1996). Following the study by Gould and Paserman (2003), existing papers also document the effects of inequality on the marriage market. However, the specific impacts of inequality on divorce have yet to be explored. In this paper, we empirically show that the interplay of inequality and gender imbalance could increase divorce rates, which can be considered a form of socioeconomic instability.[15]

2 Data and Descriptive Patterns

The main analyses involve regressing divorce on sex ratio, controlling for various control variables and fixed effects. We also conduct 2SLS estimations using several instrumental variables for robustness. Below, we describe the datasets and how we construct the main variables. Our empirical analyses draw on several datasets. Provincial aggregate results are derived from provincial panel data, while individual-level analyses combine census and survey data. The different datasets provide different measures of divorce, enabling us to get a comprehensive understanding of how sex imbalance affects marriage stability.

[14] Zhang (2017) is pertinent to our study as we both investigate the effects of the One Child Policy on divorce rates in China. However, our focus diverges as we examine different affected populations and propose distinct mechanisms. Zhang (2017) explores the impact of reduced fertility on mothers' divorce decisions, while our research examines how the biased gender ratio experienced by the birth control generation affects their own divorce behaviors.

[15] Edlund and Pande (2002) document that the divorce rate even shapes the political gender gap in the United States, and they provide evidence from longitudinal data indicating that following divorce, women are more likely to support the Democratic Party.

2.1 Provincial Panel Data

We begin with province-level analyses using data from the China Statistical Yearbook between 1997 and 2020. The National Bureau of Statistics (NBS) compiles these yearbooks, which provide comprehensive annual data for each province. The data includes the number of divorce cases, demographic composition, and social economics characteristics such as GDP per capita. In addition, we use 2000 census sample data to infer the sex ratio in the marriage market for each province each year. For example, we use the sex ratio of those aged 15-35 in the 2000 Census sample as a proxy for the sex ratio of the population aged 20-40 in 2005.

For the provincial analyses, the main outcome variable is the number of divorce cases during a given year per 1000 people aged 15-64. For robustness, we employ alternative measures of provincial divorce rates, such as the divorce-to-marriage ratio. We measure the sex ratio in the marriage market by the male-to-female ratio of the population aged 20-40. More than 90% of women who divorced got divorces when they were of this age range.[16] The descriptive statistics of the key variables are shown in Appendix Table B.1 Panel A.

2.2 Census Data

The second data source is the individual-level census sample data for 2000, 2005, 2010, and 2015. The census sample data provides information on each woman's current marital status, prefecture of residence, and other demographic and socioeconomic characteristics such as fertility, education attainment, and employment.

We focus our analyses on the currently married or divorced women aged 20-40. We measure the sex ratio following the previous literature (Brainerd, 2017; Ogasawara and Igarashi, 2021). For women living in prefecture c born in year y, the sex ratio is constructed in the following way:

$$Sex\_ratio_{cy} = \frac{\sum_{t=y-8}^{y+2} \# \, of \; men \; born \; in \; year \; t \; in \; prefecture \; c}{\sum_{t=y-8}^{y+2} \# \, of \; women \; born \; in \; year \; t \; in \; prefecture \; c} \tag{1}$$

To avoid the endogeneity issue caused by gender-biased local labor market shocks, which can affect both the gender composition of the labor force and people's marriage

[16]The age of divorce data are from the 2010 wave of the China Family Panel Studies (CFPS), where the marriage history information is available. A detailed description of the CFPS data is in Section 2.3.

behavior, we use the 2000 census data to construct this measure.[17] We restrict the birth year gap to [-8,+2] as women tend to marry men who are older than themselves.[18] Additionally, we consider prefectures as a reasonable range for the marriage market for most people.[19] The sex ratio is then merged with 2005, 2010, and 2015 census data based on each woman's prefecture of residence and birth year.

Our analyses use census data in 2005, 2010, and 2015. The divorce rates were very low before 2005. A comprehensive revision of the marriage law took effect in 2001, aiming to enhance women's rights after divorce. It was after the revision that the divorce rate began to rise. Hence, data from the census years after 2001 are more relevant for our study, which focuses on the increasing divorce rates since the early 2000s.

The key variables are summarized in Appendix Table B.1 Panel B. The natural sex ratio at birth is typically around 1.05 or 1.06 (Sen, 1992). However, the sex ratio of the marriage-age population tends to be lower due to various factors, including natural factors,[20] war casualties, aging, etc. The average male-to-female ratio for the population aged 20-40 in the United States was approximately 1.01 in 2010.[21] In our census sample data, which covers 2005, 2010, and 2015, the sex ratio is around 1.03. However, it's important to note that this figure includes individuals born before and after the birth control policy. When we consider only those born before the one-child policy, the average sex ratio is 1.01, while for those born after, it rises to 1.08. The disparity highlights the impact of the birth control policy on the sex ratio.

2.3 CFPS 2010

The third dataset is the 2010 wave of the China Family Panel Studies (CFPS). This longitudinal survey is conducted by the Institute of Social Science Survey at Peking University and provides comprehensive information on households and individuals. The CFPS sample encompasses 25 provinces in China, which covers approximately

[17] The 2000 census sample used in this paper also has the largest sample size compared to 2005, 2010, and 2015 census waves, which makes the calculation of sex ratio for each $prefecture \times birth\ year$ cell more precise. We have tried measuring the sex ratio using all the census waves, and the results remained similar. The estimation results are reported in Appendix Table B.2.

[18] As some prefecture-year cells have few observations, to avoid the impacts of outliers, we Winsorize sex ratios at the 1st and 99th percentiles. The results obtained without Winsorizing are quantitatively similar. We have also tried sex ratios constructed using other age gap ranges; the results are similar. The corresponding estimation results are reported in Appendix Table B.2.

[19] For rural areas, we conduct robustness tests using county-level sex ratios. The results remain robust. The results are shown in Appendix Table B.3.

[20] Men often exhibit higher susceptibility to different health conditions compared to women, referred to as "male disadvantage". The disadvantage can be attributed to biological differences, such as variations in maturity, sex chromosomes, and hormones (Ritchie and Roser, 2019).

[21] The number is derived from the authors' calculation using data from United Nations and Social Affairs (2022).

94.5% of the total Chinese population.[22]

The survey data possesses two key strengths, making it a valuable complement to the main analyses using the census data. Firstly, it provides rich information on the complete marriage history of women, enabling the construction of an indicator for "ever divorced" instead of relying solely on the current marital status. Specifically, the marriage history includes details such as the year of each marriage, the year of each divorce, and the total number of marriages and divorces. Leveraging this data, we generate an indicator that identifies individuals who have ever experienced divorce, serving as an additional proxy for divorce behaviors. Secondly, the survey data collects rich information on household and individual characteristics such as the number of siblings, income, psychological well-being, etc. The rich information allows us to account for additional individual characteristics, further addressing potential biases related to omitted variables.

For the sex ratio, we use the same data as in the main analyses of the census data. The sex ratio is merged with the survey data based on each woman's prefecture of residence and birth year. The key variables are summarized in Appendix Table B.1 Panel C.

3 Main results

3.1 Gender Imbalance and Divorce Rate: Provincial Panel Data

We first report the regression results at the province level. Specifically, for each province p in year t, we estimate the following equation:

$$Divorce\_rate_{pt} = \beta_0 + \beta_1 Sexratio20\_40_{pt} + X_{pt}\eta + \gamma_p + \mu_t + \epsilon_{pt} \qquad (2)$$

where $Sexratio20\_40_{pt}$ is the male-to-female ratio of population aged 20-40 in year t in province p. X_{pt} is a vector of control variables, including GDP per capita, average household size, and average years of education. γ_p and μ_t are province fixed effects and year fixed effects. With only 31 provinces in mainland China, there is a potential concern that there may not be enough clusters for the asymptotics (Donald and Lang, 2007). To address this issue, we conduct the wild cluster bootstrap test, as suggested by Cameron et al. (2008), and report the p-values in square brackets.

The results are shown in Table 1. In all regressions, the coefficients of the sex ratio variable are positive and significant, which suggests a positive relationship between the sex ratio and divorce rate. The magnitudes of the coefficient are non-negligible. Taking

[22]The excluded provinces/cities/autonomous regions in mainland China are Xinjiang, Tibet, Qinghai, Inner Mongolia, Ningxia, and Hainan.

the coefficient of Column (4) as an example, the coefficient indicates that a one-standard-deviation increase in the sex ratio can increase the divorce rate by 11.3%.[23] The results remain robust if the outcome variable is the number of divorces divided by the number of marriages in the province for a given year. This measure addresses the concern that there are more divorces because more people are married.

Table 1: Gender Imbalance and Divorce Rate at Provincial Level

	Divorce Cases Divided by Population Aged 15-64				Divorce Cases Divided by Marriage Cases
	(1)	(2)	(3)	(4)	(5)
Sexratio20_40	8.114**	6.212**	5.504*	5.516*	57.099*
	(2.953)	(3.013)	(2.906)	(3.018)	(28.999)
	[0.012]	[0.060]	[0.088]	[0.097]	[0.065]
ln(GDP per capita)		0.838	0.891	0.885	-11.906**
		(0.556)	(0.553)	(0.563)	(5.164)
Household size			0.340*	0.342*	9.582***
			(0.172)	(0.194)	(1.945)
Years of education				0.015	-0.986
				(0.288)	(1.642)
Obs.	744	744	744	744	651
Adj. R-sq	0.878	0.880	0.880	0.880	0.869
Mean of Dep. Var.	2.503	2.503	2.503	2.503	28.449

NOTE: Standard errors clustered at province level are reported in parentheses. The entries in square brackets are $p - values$ from the wild cluster bootstrap test with the null of the coefficient equal to zero. ***p<0.01, **p<0.05, *p<0.1.

3.2 Gender Imbalance and Women's Probability of Divorce: Census Data

The province-level analyses are not merely aggregates of the individual-level analyses. The provincial data is annual and thus more frequent, and the measurement of divorce is different. On the other hand, provincial analyses can be susceptible to omitted variable bias, and provinces can be too large when it comes to marriage markets. Therefore, we turn to individual-level analyses, where more factors can be controlled for, and the

[23] In the dataset, the standard deviation of sex ratio is .051, and the mean divorce rate is 2.503. Therefore, with a coefficient of 5.516, a one-standard-deviation increase in the sex ratio can increase the divorce rate by $5.516 \times .051/2.503 \times 100\% \approx 11.3\%$

marriage market can be analyzed at the prefecture level. Specifically, for each woman i living in prefecture c born in year y and surveyed in census wave w, we estimate the following equation:

$$Divorce_{icyw} = \alpha_0 + \alpha_1 Sex\_ratio_{cy} + Z_{icyw}\eta + \delta_{cw} + \kappa_{yw} + \epsilon_{icyw} \qquad (3)$$

where Z_{icyw} is a vector of controls including whether the woman has any son, her ethnicity (whether she identifies as Han), her employment status, whether she is a migrant, whether she lives in an urban area, as well as fixed effects for education level and number of children. δ_{cw} are the prefecture times census wave fixed effects and κ_{yw} are the birth year times census wave fixed effects. Standard errors are clustered at the prefecture level.

Before diving into the analyses, we first show that there are significant variations in sex ratios across prefectures and birth cohorts, given the variation in policy enforcement stringency. Figure 1 shows the residualized sex ratios obtained by regressing sex ratios on prefecture fixed effects and birth year fixed effects. The six prefectures are picked randomly from the sample. Two key observations emerge. First, there is no consistent pattern across prefectures. The sex ratio increased and decreased at different times in different prefectures, and there is no persistent trend in most prefectures. Secondly, there is substantial variation across time in most prefectures. In terms of magnitude, a change of .05 means there are five more men per 100 women.

The main results are reported in Table 2. The coefficient of the sex ratio variable is positive and significant when pooling all the censuses. Specifically, a coefficient of 0.020 indicates that a one-standard-deviation increase in the sex ratio corresponds to a 9.6 percent increase in the probability of being currently divorced.[24] The magnitude is close to what has been found in previous literature to be the impact of unilateral divorce laws on the divorce rate in the United States (Friedberg, 1998; Wolfers, 2006; Drewianka, 2008). When analyzing different census waves separately, we find that the coefficient of sex ratio is positive and marginally significant for the 2005 census, while it becomes larger and more significant for the two subsequent waves. This disparity is likely due to the rarity of divorce in the early 2000s. The increase in divorce rates over time enlarges the variation in divorce rates across prefectures and cohorts, making the impact of the sex ratio on divorce more apparent and significant in subsequent censuses.

We conduct a series of robustness tests. All the results are reported in Appendix Table B.2. First, we alter the age range to 25-45 to check the sensitivity of the results to

[24] In our dataset, the standard deviation of the sex ratio is 0.077, and on average, 16 out of 1000 women are currently divorced. Therefore, a one-standard-deviation increase in the sex ratio is associated with a $0.077 \times 0.020/0.016 \times 100\% = 9.6\%$ increase in the likelihood of being currently divorced.

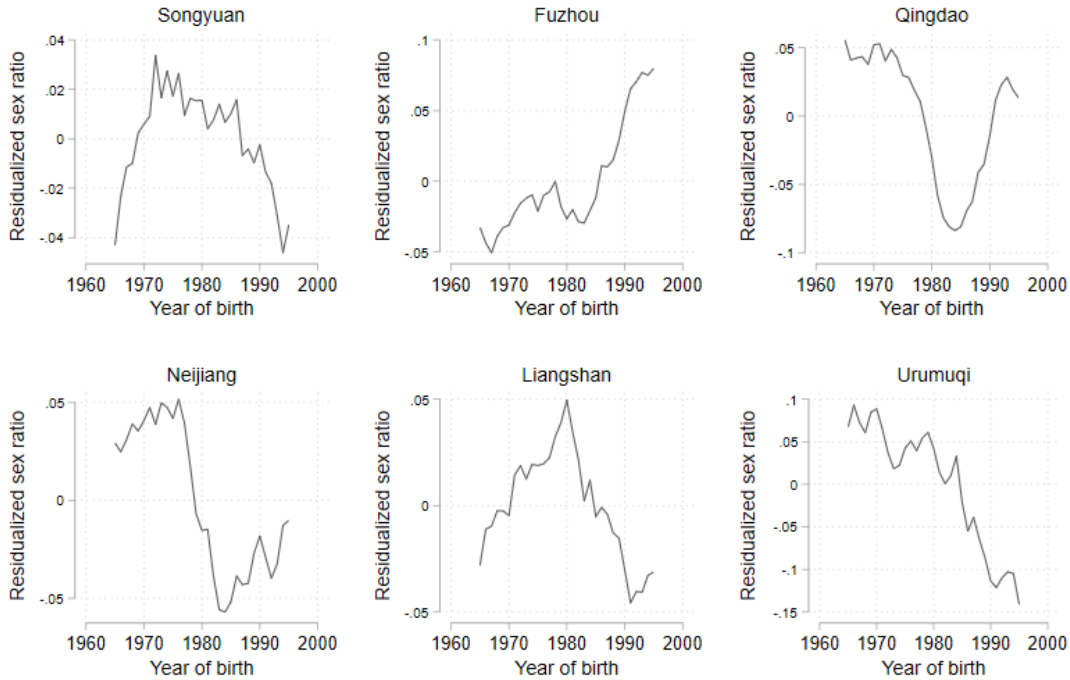

Figure 1: Residualized Sex Ratios

NOTE: The figure shows the residualized sex ratios obtained from regressing sex ratio on prefecture fixed effects and birth year fixed effects.

the specific age range. Secondly, we changed the age cohort for sex ratio calculation to [-6,+1] or [-8,+2]/[-7,+3].[25] Thirdly, we used the census wave-specific sex ratio instead of the sex ratio in the base year 2000. These robustness tests provide additional support to the positive relationship between the sex ratio and the probability of being currently divorced.

We take various measures to address potential threats to the identification. Firstly, the impact of the sex ratio on women's education attainment and labor market outcomes could affect women's independence and subsequently influence the probability of divorce. In the main analyses, we have controlled for the education and employment status of the women. The results remain robust when we further include industry and occupation fixed effects, which suggests that the potential impact of the increased independence is not a main concern for our findings. Secondly, the changes in social attitudes towards

[25] [-6,+1] refers to $sexratio_{cy} = \frac{\sum_{t=y-6}^{y+1} \# \; of \; men \; born \; in \; year \; t \; in \; prefecture \; c}{\sum_{t=y-6}^{y+1} \# \; of \; women \; born \; in \; year \; t \; in \; prefecture \; c}$, and [-8,+2]/[-7,+3] refers to $sexratio_{cy} = \frac{\sum_{t=y-8}^{y+2} \# \; of \; men \; born \; in \; year \; t \; in \; prefecture \; c}{\sum_{t=y-7}^{y+3} \# \; of \; women \; born \; in \; year \; t \; in \; prefecture \; c}$

Table 2: Gender Imbalance and Women's Probability of Divorce: Censuses

	Pooled (1)	Census 2005 (2)	Census 2010 (3)	Census 2015 (4)	IV (5)
Sex ratio	0.020***	0.010*	0.023***	0.020***	0.099**
	(0.004)	(0.006)	(0.005)	(0.007)	(0.041)
Han ethnicity	-0.003***	-0.005***	-0.003**	-0.001	-0.002
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)
Having a son	-0.005***	-0.004***	-0.006***	-0.006***	-0.005***
	(0.000)	(0.000)	(0.000)	(0.001)	(0.001)
Employed	0.001**	-0.002**	0.001**	0.003***	-0.000
	(0.000)	(0.001)	(0.001)	(0.001)	(0.001)
Migrant	-0.003***	-0.005***	-0.002***	-0.003***	-0.004***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Urban	0.008***	0.009***	0.009***	0.005***	0.009***
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
1st-Stage F-stat					6.393
Observation	1,150,344	339,405	593,688	217,250	461,665
Adj. R-sq	0.018	0.017	0.018	0.016	0.000
Mean of Dep. Var.	0.016	0.011	0.018	0.020	0.018

NOTE: Prefecture × wave fixed effects and birth year × wave fixed effects, education, and number of children fixed effects are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

divorce should not account for our findings as the results remain robust when we control for province times birth year fixed effects. Thirdly, as documented by Chen et al. (2021) and Alm et al. (2022), people in some regions resort to "fake divorces" to get additional quotas for house purchasing. As the previous literature has established that a biased sex ratio can increase demand for houses, the observed positive impact of sex ratio on divorce can be caused by this impact of sex ratio on house demand, with the increased divorce rate serving as a method to secure quotas for purchasing houses. To address the potential effects of fake divorces, we control for the house purchasing quota policy, and the results remain robust. Finally, as we can only observe the sex ratio of the prefecture of residence but not the sex ratio the women were exposed to when they divorced, it is possible that migrants were exposed to a different sex ratio when making divorce decisions. To address this concern, we drop the migrants, and the results remain. All the results are reported in Appendix Table B.3.

Given the extensive time span of over twenty years, it is unlikely that parents engaged in sex selection practices with the anticipation of future conditions in the marriage market. Therefore, we consider the potential omitted variable problem the largest threat to our results. In Section 3.4, we explore and rule out the major potential confounding factors. It is important to note that our sex ratio measure is highly granular, and we have controlled for prefecture fixed effects and birth year fixed effects. By doing so, any remaining threats to identification should vary at the prefecture times birth year level.

Finally, we experiment with multiple instruments to address concerns about potential omitted unobservables. However, the high granularity of the sex ratio measure results in weak first stages, causing the second-stage estimations to be 1-5 times the size of the OLS estimations and less precisely estimated. To mitigate potential issues related to the instruments, our primary analyses focus on Ordinary Least Squares (OLS) estimates. The preferred instrumental variable result is included in Table 2, Column (5). Specifically, the instrumental variable is the interaction of a measure of prefecture-level gender norms and the indicator for being born in and after 1979, the starting year of the One-Child Policy.[26] The coefficient on sex ratio remains positive and significant in our 2SLS estimation. See Appendix Section B.3 for results of other instrumental variables and detailed discussions.[27]

3.3 Gender Imbalance and Women's Probability of Divorce: Individual Survey Data

The census data only has information on the current marital status of women. If a woman divorced and remarried, they will be coded as "married" rather than "divorced." This will lead to an underestimation of the impact of sex ratio on divorce. To address this issue, we exploit the CFPS data where the whole marriage history is observed. Specifically, we estimate the following equation:

$$Ever\_divorced_{icy} = \alpha_0 + \alpha_1 Sex\_ratio_{cy} + Z_{icy}\eta + \delta_c + \kappa_y + \epsilon_{icy} \qquad (4)$$

[26]The gender norm variable is constructed using the 2010 wave of the Chinese General Social Survey (CGSS) dataset. Specifically, each interviewee responded to 5 statements related to gender norms, providing a score ranging from 1 to 5 to indicate their levels of agreement. A higher score reflects a greater level of agreement. The five statements are: "Women prioritize their families, while men prioritize their careers," "Men are inherently stronger than women in terms of abilities," "A successful marriage is preferable to a successful career," "When the economy is struggling, female employees should be laid off first," "Household chores should be divided equally between spouses." We recode the score of the last question so that a higher average score of the five questions reflects a higher degree of gender discrimination. We expect a positive sign in the first stage, given that sex selection, following the launch of the One-Child Policy, is likely to be more prominent in regions characterized by stronger gender discrimination.

[27]We only show results that have a significant first-stage coefficient. We have experimented with multiple instruments, incorporating insights from Zhang (2017), such as minority fractions across cohorts, excess fertility residuals, and fine rates. All of them result in weak first stages.

where Z_{icy} is a vector of individual characteristics. In addition to the control variables in the regressions using census data, we control for the indicator for co-habitation before marriage and fixed effects for the number of male and female siblings. The standard errors are clustered at the prefecture level.

The regression results are reported in Table 3. Even after controlling for a rich set of control variables and fixed effects, the coefficients of sex ratio remain positive and significant, implying that an increase in sex ratio will increase the probability of a woman experiencing divorce. Specifically, a coefficient of 0.098 indicates that a one-standard-deviation increase in the sex ratio corresponds to a 23 percent increase in the probability of ever getting a divorce.[28] Compared to the estimation using the census data, the larger magnitude of this effect aligns with the recognition that the census data tends to underestimate, as about half of the divorced women remarry.[29] The results remain robust when we control for the women's income.

Table 3: Gender Imbalance and Women's Probability of Divorce: CFPS 2010

1{Ever divorced}	(1)	(2)	(3)	(4)
Sex Ratio	0.085*	0.098**	0.098**	0.098**
	(0.047)	(0.046)	(0.047)	(0.046)
Controls	-	Y	Y	Y
Male and female siblings FE	-	-	Y	Y
Income	-	-	-	Y
Obs.	8,799	8,798	8,717	8,715
Adj. R-sq	0.019	0.073	0.074	0.075
Mean of Dep. Var.	0.028	0.028	0.028	0.028

NOTE: Fixed effects for prefecture, birth year, education, and the number of children are controlled for all columns. All the control variables for the census regressions are controlled for in Columns (2)-(4). Additionally, the indicator for cohabitation before marriage is controlled. Column (4) controls for the inverse hyperbolic sine of income. If controlling for log income instead, the coefficient increases slightly with no change in statistical significance, and the sample size shrinks to three quarters. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

We address one potential omitted variable bias by controlling for sibling fixed effects. Specifically, the birth control policy can simultaneously influence the sex ratio, the sex composition of siblings, and the number of siblings. There has been literature on the "little prince/princess" phenomenon, which suggests that the children of the birth

[28]One standard deviation in sex ratio is 0.065, and the mean divorce rate is 0.028. Therefore, a one-standard-deviation increase in the sex ratio corresponds to $0.065 \times 0.098/0.028 * 100\% \approx 23\%$

[29]The fraction of remarried women among those who have ever divorced is calculated using the 2005 census and CFPS 2010 data, where the relevant information is available.

control generation can be spoiled as they gain seemingly excessive amounts of attention from their parents and grandparents (Cameron et al., 2013). If the spoiling changes the children's expectations of marriage, their divorce decisions can also be affected. However, as shown in Column (3), the positive impact of sex ratio on divorce remains robust when the sibling fixed effects are controlled for. Additionally, the magnitude of the coefficient changes little, indicating that the impact of the lack of siblings is not a significant contributing factor to the observed relationship between gender imbalance and divorce rates.

3.4 Rule Out Other Potential Explanations

Previous literature and anecdotal evidence suggest three potential channels that may play a role in the observed impacts of sex ratio on divorce. In this section, we examine these alternative explanations.

First, a higher sex ratio may lead women to marry earlier as they may encounter more men when they are young. Marrying younger may result in a higher probability of divorce as younger wives' ability to deal with challenges in marriage can be low (Rotz, 2016). To check if this is a relevant story, we control for the fixed effects of age at first marriage in addition to the baseline specification. The result is shown in Table 4 Column (2). Column (1) shows the baseline results copied from Table 2. The coefficient of sex ratio in Column (2) remains positive and significant, and the magnitude is very close to the baseline result. The result suggests that changes in the ages at the first marriage do not explain the impact of sex ratio on divorce. Further analyses indicate that an increase in the sex ratio results in a very slight increase in the marriage age rather than a decrease.

Secondly, besides affecting sex ratios, birth control policies also contribute to the prevalence of many one-child households. In the Chinese context, especially in rural areas, it is common for married couples to live with the husband's parents. A higher likelihood of the husband being the only child results in a higher probability of women co-living with their parents-in-law, which may be associated with an increased divorce rate.[30] Therefore, co-living can lead to omitted variable bias in our setting. To address this concern, we calculate the co-residence fraction at the prefecture times birth year level

[30] Co-residence with parents-in-law may increase the divorce rate for several reasons. Firstly, daughters-in-law and mothers-in-law may face conflicts over household decision-making, leading to tension and stress within the family. These conflicts can arise from differing opinions on financial matters, child-rearing practices, and daily routines. Secondly, living together means that the daughter-in-law must take on the responsibility of caring for her in-laws, which can be particularly demanding when they are ill or require constant attention. This added burden can strain the marital relationship, as the daughter-in-law may feel overwhelmed and unsupported. Additionally, the lack of privacy and autonomy in a multi-generational household can further contribute to marital dissatisfaction and conflict.

Table 4: Rule Out Other Explanations: Censuses

	Baseline	Marriage age	Fraction of wives co-living with husband's parent	Gender Edu Gap	Add all controls
	(1)	(2)	(3)	(4)	(5)
Sex ratio	0.020***	0.021***	0.021***	0.020***	0.022***
	(0.004)	(0.005)	(0.008)	(0.004)	(0.008)
Co-live			0.022***		0.025***
			(0.003)		(0.003)
Education gap				-0.004***	-0.004**
				(0.001)	(0.002)
Marriage age FE	-	Y	-	-	Y
Obs.	1,150,344	933,093	795,372	1,150,344	695,204
Adj. R-sq	0.018	0.019	0.019	0.018	0.020
Mean of Dep. Var.	0.016	0.016	0.017	0.016	0.016

NOTE: Fraction of wives co-living with husband's parent is at the prefecture times birth year level for each census wave, distinguishing between urban and rural areas. The gender education gap is calculated in the following way. For a woman observed in census wave w living in prefecture c born in year y, we first calculate the average education gap between men and women for each census wave × prefecture × birth year cell. Then, we take the average of the gaps for birth years in the range $[y-8, y+2]$. This gap is taken as a proxy for the education gap the woman faced in her marriage. Prefecture times wave, birth year times wave, education, and number of children fixed effects are controlled for all columns. Column (1) shows the baseline results copied from Table 2. All the control variables for the census regressions are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

for each census wave, distinguishing between urban and rural areas. We then control for the co-living fraction in the regression. The result is shown in Table 4, Column (3).[31] The coefficient of sex ratio remains positive and significant. This indicates that the omitted variable bias caused by co-living is not a primary concern in our context. The coefficient of the co-living fraction is positive and significant, suggesting that co-living with parents-in-law increases the probability of divorce for women.

Thirdly, birth control policies can increase the education attainment of both men and women, with a potentially larger impact on women's education, leading to a reduction in the gender education gap (Chae et al., 2023). This reduced gap may also influence the probability of divorce, introducing an omitted variable bias. To address this concern, we control for the education gap. Specifically, for a woman observed in census wave w

[31] The observations decrease because we exclude the prefecture × birth year × urban cells with too few observations in order to improve the precision of the co-living fraction.

living in prefecture c born in year y, we first calculate the average education gap between men and women for each census wave times prefecture times birth year cell. Then, we take the average of the gaps for birth years in the range $[y-8, y+2]$ and use it as a proxy for the education gap the women faced in her marriage. The result is shown in Table 4, Column (4). In the existing literature, whether a narrowed gender education gap destabilizes marriages remains controversial (see Bertrand et al. (2015) and Binder and Lam (2022)). In any case, it does not affect the main results.

Finally, when we include all controls in the same regression, the coefficient for the sex ratio is positive and significant and has a similar magnitude as the main results.

There is another possible explanation. The increased mating competition due to gender imbalances might motivate men to enhance their work efforts, such as extending work hours or migrating for jobs, as suggested by Wei and Zhang (2011b) and Zhang et al. (2021). However, if these adaptations render them less present in personal relationships—for example, if they are unable to spend time with their spouses or struggle with relationship maintenance—this could also lead to higher divorce rates. We address this hypothesis in two ways. Firstly, as elaborated in the subsequent section, our findings indicate that women report greater satisfaction in their marriages. This contradicts the assumption that women would be less content if their husbands were frequently absent or inept in handling relationships. Secondly, we calculate the fraction of migrant workers among males at census year times prefecture times birth year level. Controlling for this fraction does not change our results.

4 Mechanism

Overall, we observe a robust pattern that facing a higher sex ratio will increase women's probability of divorce. We investigate potential channels that may contribute to this pattern to understand this relationship further.

As mentioned before, the impact of sex ratio on divorce is ex-ante ambiguous. On the one hand, a higher sex ratio gives women higher bargaining power in the marriage market. As a result, they can marry a better husband, which should decrease the probability of divorce. On the other hand, a higher sex ratio also leads to more outside options, which can make marriages less stable. We provide a simple model in Appendix Section A to illustrate the effects of the two forces. The model demonstrates that, depending on different conditions, the same theoretical framework can yield either positive or negative overall impacts of sex imbalance on divorce rates. In our subsequent analyses, we thoroughly examine both of these directions.

4.1 Do Women Marry Up When the Sex Ratio is Higher?

We first examine whether women experience marriages of higher quality in the presence of a higher sex ratio. The implications of the main results are different if there is excessive friction in the marriage market that prevents women from finding a preferred husband.[32] To examine whether women marry up when faced with a higher sex ratio, we regress various husband quality measures on sex ratio. The results are reported in Table 5 Column (1)-(3). Only couples married within three years are included in the sample to limit survivorship bias. The results indicate that women faced with a higher sex ratio tend to marry men with more years of education and higher monthly income. Women with rural *Hukou* are more likely to marry husbands with urban *Hukou*.[33]

Moreover, given the important role of houses in marriage, we examine the impacts of sex ratio on housing quality. Specifically, in the Chinese context, where parents-in-law often make substantial contributions to marriage-related house purchases, these measures can serve as proxies of parental affluence. Column (4)-(6) of Table 5 show the regression results. As expected, a higher sex ratio is associated with an increase in housing quality measured by house area, number of rooms, and house price.

Other than social and economic characteristics, we recognize that unobserved features, such as personality, also play a role in determining the quality of marriage. While lacking information on the unobserved features, we can examine the quality of marriage directly by examining women's happiness in their marital relationships. Specifically, we examine whether women are less likely to experience mental health issues in marriage when faced with a higher sex ratio. To do this, we utilize the CFPS data, which has information on nonspecific psychological distress collected using the Kessler 6 Scale (K6) instrument (Kessler et al., 2002). During the survey, the interviewees were asked about the frequency of feeling "upset", "stressed", "cannot calm down", "hopeless", "everything is difficult" and "life is meaningless". The responses were coded into indexes ranging from 0 to 4: 0 for "none of the time," 1 for "a little of the time," 2 for "some of the time," 3 for "most of the time," and 4 for "all of the time."

We regress these indexes on the sex ratio for currently married women, and the results are shown in Table 6. All the coefficients of sex ratio are negative, which indicates that a higher sex ratio decreases the probability of the wife feeling depressed in various ways.

[32] It is also theoretically possible that, knowing there is a high likelihood of a second chance, women may be less stringent about their first marriage. Therefore, a higher sex ratio does not necessarily lead to higher marriage quality.

[33] Compared to their rural counterparts, urban *Hukou* holders enjoy superior social security benefits, including better healthcare, pensions, unemployment benefits, and minimum living standards.

Table 5: Gender Imbalance and Husband Quality

	Husband Characteristics			Housing Quality		
	Education	log Monthly Income	Has Urban Hukou	lnArea	# of Rooms	lnHouseprice
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio	0.585**	0.277**	0.105***	0.779***	1.175***	0.298*
	(0.236)	(0.113)	(0.039)	(0.174)	(0.275)	(0.156)
Obs.	103,264	28,549	72,238	103,188	103,188	24,226
Adj. R-sq	0.627	0.536	0.152	0.312	0.256	0.575
Mean of Dep. Var.	10.684	6.483	0.083	4.376	3.200	10.227

NOTE: Only couples married within three years are included in the sample to limit survivorship bias. As Census 2015 does not have information on marriage year, only Census 2005 and 2010 are used. Columns (2) and (6) use Census 2005 only due to data availability. Column (3) only includes women who have rural *Hukou*. Prefecture × wave fixed effects, birth year × wave fixed effects, and wife's education fixed effects are controlled for all columns. Fixed effects for the year of construction of the house are controlled for Columns (4)-(6). All the control variables for the census regressions are controlled for all columns except for the number of children or the indicator for having a male child. These two variables are omitted as fertility occurs after marriage and is not expected to impact husband and housing characteristics directly. Results are similar if they are included in the regressions. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

The results remain robust when we control for the income of the women.[34]

In conclusion, we observe that women marry better when exposed to a higher sex ratio, which could potentially decrease the probability of divorce. The results suggest that the higher divorce rate cannot be solely attributed to the frictional nature of the marriage market, where women struggle to find suitable partners in their first marriage. Instead, the results suggest that another force, which we identify as the outside option channel, plays a dominant role in increasing divorce rates.

4.2 Outside Option

In this section, we provide evidence for the outside option channel. This channel suggests that when faced with gender imbalances, individuals can have more alternative options, making marriages less stable and leading to a higher likelihood of divorce.

[34] Additionally, it has been documented in the previous literature that women's labor supply can decrease when they have higher within-household bargaining power. We verify this result using our data and find that when faced with a higher sex ratio, women decrease labor supply both at the intensive and extensive margins. Moreover, we find that women do fewer house chores during weekdays if they are faced with a higher sex ratio. These results suggest that women have higher bargaining power within household when sex ratio increases.

Table 6: Gender Imbalance and Wife's Mental Health

The frequency of feeling	Upset	Stressed	Cannot calm down	Hopeless	Everything is difficult	Life is meaningless	Overall depression
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sex ratio	-0.624***	-0.117	-0.425**	-0.256*	-0.448**	-0.207	-0.957**
	(0.236)	(0.197)	(0.186)	(0.152)	(0.203)	(0.173)	(0.449)
Obs.	8,558	8,565	8,564	8,545	8,565	8,558	8,530
Adj. R-sq	0.056	0.046	0.056	0.051	0.082	0.060	0.088
Mean of Dep. Var.	0.763	0.569	0.493	0.325	0.562	0.333	-0.014

NOTE: This table shows estimations using the CFPS data. Only women who are currently married are included in the sample. Outcomes of Columns (1)-(6) are indexes that reflect the frequency of feeling depressed in various ways. These indexes range from 0 to 4: 0 for "none of the time", 1 for "a little of the time", 2 for "some of the time", 3 for "most of the time", and 4 for "all of the time." The outcome of Column (7) is the Principal Component Analysis (PCA) scores of the first principal component. Prefecture, birth year, education, and number of children fixed effects are controlled for all columns. All the control variables for the census regressions and the indicator for cohabitation before marriage are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

4.2.1 Do Women Remarry More Easily after Divorce?

We first check if divorced women are more likely to remarry if faced with a higher sex ratio in the local marriage market. If women are getting divorced because they no longer have the demand for marriage, we would not observe the impact of the sex ratio on remarriage. In other words, if we do find an effect, it suggests that the availability of alternative options may play a role in women's decision to divorce. The anticipation of easy remarriage could reduce women's reluctance to divorce, thereby contributing to the higher divorce rates of women faced with higher gender imbalances.

Table 7 reports our estimation results on the impacts of sex ratio on remarriage using CFPS data. We examine both ever-divorced and ever-widowed women samples, as women can remarry after divorce or during widowhood. For both samples, we find that women who have experienced the loss of a partner are more likely to get remarried when confronted with higher sex ratios. Overall, the results suggest that women can find a new partner relatively more easily when faced with higher sex ratios, which may increase the likelihood of divorce.

4.2.2 Is the Quality of the Subsequent Husband Better for Remarried Women?

We then investigate whether a higher sex ratio increases the quality of the subsequent partner for remarried women. If this impact is observed, it suggests that the availability of alternative options may play a role in women's decision to divorce. If women expect

Table 7: Gender Imbalance and Remarriage

1{Ever-remarry}	Divorced	Divorced & Widowed
	(1)	(2)
Sex ratio	3.267**	2.550***
	(1.393)	(0.933)
Obs.	201	400
Adj. R-sq	0.258	0.139
Mean of Dep. Var.	0.507	0.450

NOTE: Prefecture, birth year, education, number of children, male and female siblings fixed effects are controlled for all columns. All the control variables for the census regressions and the indicator for cohabitation before marriage are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

that potential new partners have reasonably high quality and can lead to a more satisfying and stable future relationship, they are more likely to seek divorce.

Table 8 shows the results. The sample includes all currently married or remarried women. The coefficient for the remarriage indicator implies that the new partners of remarried women have fewer years of education and lower monthly income. On the other hand, for women with rural *Hukou*, the subsequent partners are more likely to have urban *Hukou*. The results suggest that in most aspects, the husbands of remarried women are of lower quality than those of women who remain in their first marriage. There are at least two potential reasons. Firstly, women who once divorced are less preferred for men, so only men with lower quality marry divorced women. Secondly, men of high quality were kept in the first marriage. Therefore, the husbands' quality of divorced women can change from very bad to not so good but cannot be better than those who remain in their first marriage.

Our results suggest that a higher sex ratio can mitigate the reduction in husband's quality. Specifically, women who face a higher sex ratio are more likely to remarry a husband with higher education than women who face a lower sex ratio. For women with rural *Hukou*, a higher sex ratio further increases the likelihood of remarrying a husband with urban *Hukou*.

The magnitudes of the coefficients for the remarriage indicator and the interaction term suggest that, at least in terms of education and income, the impact of the sex ratio is not substantial enough to offset the costs associated with remarriage. In other words, remarriages are not better than first marriages, even with the highest sex ratio. We interpret this as indicating that, for many women, divorce serves as an escape from

Table 8: Gender Imbalance and Quality of the Subsequent Husband

Husband's	Year of Education (1)	log Monthly Income (2)	Has Urban Hukou (3)
Sex ratio_demeaned	-0.120	0.103	0.026
	(0.147)	(0.068)	(0.020)
1{remarry}	-0.337***	-0.050***	0.042***
	(0.034)	(0.012)	(0.009)
Sex ratio_demeaned	1.057**	0.134	0.305**
×1{remarry}	(0.531)	(0.192)	(0.146)
Obs.	207,826	197,690	149,283
Adj. R-sq	0.534	0.471	0.081
Mean of Dep. Var.	9.452	6.370	0.053

NOTE: Use census 2005 data. Column (3) uses the subsample of women with rural *Hukou*. Prefecture, birth year, and education fixed effects are controlled for all columns. All the control variables for the census regressions are controlled for all columns except for the number of children or the indicator for having a male child. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

horrible marriages, such as those with abusive husbands, rather than a means to climb the economic ladder.[35]

4.3 Heterogeneity: Inequality

Finally, we ask if women are more likely to divorce when there are more better alternative options. Specifically, we investigate if the sex ratio increases divorce more in regions with higher inequality. The intuition is that a higher sex ratio increases the number of outside options, and higher inequality increases the fraction of these outside options of higher quality than the current husband plus a divorce cost. If it is easier for women to find new partners that are better than their current husbands, the likelihood of divorce will increase. Notice that inequality may also improve the husband's quality in the first marriage, which should amplify the positive impact of sex ratio on first marriage quality and decrease the divorce rate. Therefore, the sign of the coefficient for the interaction of sex ratio and inequality is ex-ante ambiguous. A positive sign will suggest that the outside option channel dominates.

[35]In this paper, we do not necessarily view divorce as a negative event. For some women, particularly those with abusive husbands, divorce can provide a means to escape a harmful situation and improve their overall well-being. A branch of literature supports the notion that divorce may enhance women's welfare (see, for example, Stevenson and Wolfers, 2006).

Table 9: Heterogeneous Effect of Gender Imbalance: Inequality

	(1)	(2)	(3)	(4)	(5)
Sex ratio * P90/50	0.128**				
	(0.064)				
Sex ratio * P90/10		0.062*			
		(0.032)			
Sex ratio * sd(ln(income))			0.222*		
			(0.123)		
Sex ratio * sd(residual ln(income))				0.264*	
				(0.138)	
Sex ratio * Gini					0.483**
					(0.219)
Obs.	378,794	378,794	378,794	378,794	378,794
Adj. R-sq	0.017	0.017	0.017	0.017	0.017
Mean of Dep. Var.	0.017	0.017	0.017	0.017	0.017

NOTE: Inequality is calculated using household total income divided by the number of males of age 20-45 within the household (89% households have one and 8% have two such males). Inequality is calculated using the CGSS data due to data availability. Residual ln(income) is obtained by regressing log household income on fixed effects for the number of males aged 20-45 within the household and the average age of such males. Prefecture, birth year, education, and number of children fixed effects are controlled for all columns. All the control variables for the baseline census regressions are controlled for. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

Table 9 shows the results. Following Gould and Paserman (2003), we explore different inequality measures. Specifically, we use five different measures of inequality. Our approach begins by computing the household's aggregate income divided by the number of males aged 20-45 residing within the household. This construction takes into account that the entire household's income better captures the significance of parental earnings in the context of marriages in China. We take the log of this income measure and calculate its standard deviation and the Gini index for each prefecture. Additionally, we calculate the ratios between the 90th percentile and the 50th percentile of the log income measure (P90/50) as well as the ratios between the 90th percentile and the 10th percentile of the log income measure (P90/10). Moreover, inequality and social attitudes towards divorce may change simultaneously as the economy develops. To address this issue, we calculate the urban rate at census wave times province level and control for the interaction of sex ratio and the urban rate in all the regressions.

For all the inequality measures, the coefficients for the interaction of sex ratio and inequality are positive and significant, indicating that the effect of sex ratio on divorce is more pronounced in prefectures with higher inequality. The results suggest that the influence of the outside option channel outweighs the impact of the matching quality channel when considering how inequality matters for the effects of sex imbalance on divorce. Moreover, the results indicate that societies undergoing both salient inequality and substantial sex imbalance, as observed in numerous developing countries today, may face heightened concerns regarding the dissolution of marriages.

5 How about Men?

In both the analyses presented above and the model detailed in Appendix A, men are implicitly assigned a passive role in the contexts of marriage and divorce. In this section, we investigate the impacts of sex ratios on men's marriage outcomes and behaviors within the household.

We first examine the impact of sex ratios on men's divorce rates. Since same-sex marriage is not recognized in China, a divorce always involves a man and a woman. Therefore, given our robust finding that increases in sex ratio lead to higher divorce rates among women, we should also observe that increases in sex ratio similarly increase divorce rates among men. We conduct the same analyses as we do for women but using a sample of married or divorced men aged 20 to 50.[36]

Table 10: Gender Imbalance and Men's Probability of Divorce: Censuses

	Pooled	Census 2005	Census 2010	Census 2015
Divorce	(1)	(2)	(3)	(4)
Sex ratio	0.019***	0.003	0.022***	0.027***
	(0.004)	(0.005)	(0.005)	(0.006)
Obs.	1,661,579	449,980	858,222	353,377
Adj. R-sq	0.012	0.012	0.011	0.012
Mean of Dep. Var.	0.025	0.019	0.026	0.033

NOTE: Prefecture × wave fixed effects and birth year × wave fixed effects, education, and number of children fixed effects are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

Table 10 and Table 11 show the results using census and CFPS data, respectively. The coefficients on sex ratio are positive and significant across all columns, except for the year 2005 when divorces remained rare. These results suggest that higher sex ratios increase men's divorce rates.

[36]We analyze the sample of women aged 20 to 40 but use a higher upper bound of 50 for men, as a non-negligible fraction of men who have ever divorced did so between the ages of 40 and 50, while few women divorce at these ages, according to CFPS data.

Table 11: Gender Imbalance and Men's Probability of Divorce: CFPS 2010

1{Ever divorced}	(1)	(2)	(3)	(4)
Sex ratio	0.077*	0.072**	0.075**	0.076**
	(0.041)	(0.035)	(0.035)	(0.035)
Obs.	8,517	8,517	8,456	8,455
Adj. R-sq	0.018	0.075	0.076	0.076
Mean of Dep. Var.	0.043	0.043	0.043	0.043

NOTE: Fixed effects for prefecture, birth year, education, and the number of children are controlled for all columns. All the control variables for the census regressions are controlled for in Columns (2)-(4). Additionally, the indicator for cohabitation before marriage is controlled. Column (4) controls for the inverse hyperbolic sine of income. If controlling for log income instead, the coefficient barely changes, and the sample size shrinks slightly. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

We also examine the impact of sex ratios on men's remarriage rates, with results shown in Appendix Table B.5. The coefficients are all negative, suggesting that higher sex ratios decrease the probability of remarriage for men, although the statistical power is insufficient to reject the null hypothesis conclusively. The negative sign of the coefficients supports the outside option channel. Once a man gets divorced, he enters a marriage market with a higher male-to-female ratio. In these markets, the surplus of men makes it much harder for him to remarry compared to markets with a more balanced gender ratio. This increased competition in the marriage market can explain the observed negative impact on the likelihood of remarriage for divorced men.

Faced with a higher potential for divorce, what actions did men take to preserve their marriages? Based on our results and existing literature, we know that men and their parents have made substantial efforts both before and after marriage to improve women's marital satisfaction, such as by purchasing better housing assets. Our findings also indicate that women report higher levels of marital satisfaction, which may be attributed to increased efforts by men to meet their spouses' needs. In this section, we provide direct evidence that men adjust their domestic behaviors in response to higher sex ratios. Specifically, we find that men spend more time caring for family members during weekdays, as shown in Table 12.

If men have made efforts to maintain their marriages, why do we still observe increases in divorce? According to existing economic theory, specifically the Coase Theorem, if there are no transaction costs and a utility surplus exists, men should be able to transfer some of this surplus to keep women in the marriage. We provide

Table 12: Gender Imbalance and Men's Time Spent Caring for Family Members

Time spent caring for family members	Hours on weekdays	IHS(Hours) on weekdays	Hours on weekends	IHS(Hours) on weekends
	(1)	(2)	(3)	(4)
Sex ratio	0.536**	0.318**	0.206	0.200
	(0.217)	(0.132)	(0.344)	(0.167)
Obs.	8,511	8,511	8,503	8,503
Adj. R-sq	0.066	0.094	0.074	0.108
Mean of Dep. Var.	0.596	0.421	0.786	0.523

NOTE: This table shows estimations using the CFPS data. Only men who are currently married are included in the sample. Prefecture, birth year, education, and number of children fixed effects are controlled for all columns. All the control variables for the census regressions and the indicator for cohabitation before marriage are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

Table 13: Gender Imbalance and Husband's Mental Health

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
The frequency of feeling	Upset	Stressed	Cannot calm down	Hopeless	Everything is difficult	Life is meaningless	Overall depression
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Sex ratio	0.298	0.337**	0.223	-0.015	0.534**	0.037	0.699*
	(0.200)	(0.169)	(0.187)	(0.154)	(0.221)	(0.140)	(0.388)
Obs.	8,234	8,236	8,236	8,230	8,235	8,236	8,220
Adj. R-sq	0.053	0.032	0.044	0.031	0.053	0.041	0.065
Mean of Dep. Var.	0.635	0.491	0.417	0.287	0.538	0.253	-0.221

NOTE: This table shows estimations using the CFPS data. Only men who are currently married are included in the sample. Outcomes of Columns (1)-(6) are indexes that reflect the frequency of feeling depressed in various ways. These indexes range from 0 to 4: 0 for "none of the time", 1 for "a little of the time", 2 for "some of the time", 3 for "most of the time", and 4 for "all of the time." The outcome of Column (7) is the Principal Component Analysis (PCA) scores of the first principal component. Prefecture, birth year, education, and number of children fixed effects are controlled for all columns. All the control variables for the census regressions and the indicator for cohabitation before marriage are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

suggestive evidence that the utility surplus from maintaining the marriage may have been squeezed, leaving little to be transferred. Table 13 examines the impact of sex ratios on men's mental health. The estimations suggest that higher sex ratios negatively affect men's mental health, measured in various ways, indicating that men suffer mentally when sex ratios are high.

Overall, our results support our narrative. We observe increased divorce rates in the context of higher male-to-female ratios. The evidence on men's mental health suggests that they suffer more mentally when the sex ratio is higher, potentially due to their lower bargaining power at home. Faced with a high sex ratio, men's surplus from marriage has been squeezed.

6 Conclusion

While the birth control policy has been officially terminated, its profound impacts will persist in the Chinese economy for a considerable period. One of its side effects, sex imbalance, has dramatically transformed society. In addition to the impact of sex imbalance on various outcomes in marriage and labor markets documented in the previous literature, this paper sheds light on the overlooked robust relationship between sex imbalance and divorce. Further investigation reveals that the outside option channel plays a dominant role. We show that women facing a higher sex ratio are more likely to remarry after divorce, and the new partners are of higher quality compared to remarried women facing a lower sex ratio. With the expectation of easy remarriage and the prospect of better subsequent husbands, women are more inclined to divorce.

As the main mechanism we reveal is the outside option channel, regions with an imbalanced sex ratio may experience similar marriage instability. Our framework, therefore, provides a valuable perspective in other sex-imbalance settings, such as those affected by gender-biased industry structures.

We acknowledge that there could be other factors shaping the influence of sex ratio on divorce rates. Specifically, sex ratio can affect both the quality of marriages and the availability of alternative options. Factors such as social norms (for instance, whether there is bias against divorced women), legal institutions (including the extent of legal protection for women should they go through a divorce), and social support (such as the capacity of divorced women to lead independent lives and raise children on their own) can affect which force dominates, the increase in marriage quality or the increase in outside options, thereby influencing the overall impact. This variability may explain the conflicting findings in prior literature regarding the effects of male scarcity on divorce rates across different contexts. The purpose of this paper is not to establish a universally applicable rule of when the outside option channel dominates. Our study aims to offer complementary insights into a distinct scenario of female scarcity, and we hope our study can help open up new directions for future research on gender-based variations in marriage-related preferences and behaviors.

Last but not least, our findings have important implications on how policies can be designed to either slow down the trend of increasing divorce rates or address the consequences in regions with high gender imbalance and divorce rates. First, social protection programs covering the divorced population, their children, and those who cannot find mates because of gender imbalance are needed. Second, local governments may want to pay more attention to the marriage market if the regional industrial structure, immigration, or internal migration can lead to sex imbalance. Third, reducing the degree of information asymmetry in the marriage market can improve marriage-matching efficiency. For example, local governments can release more information about the relative shortage of men or women in the marriage market to guide those single adults to find the "right" destination where opportunities exist for finding both marital partners and jobs.

A A Model on Marriage, Divorce, and Re-marriage

In this section, we present a simple model to highlight how the sex ratio may impact the divorce rate. The impact is ambiguous. On the one hand, a higher sex ratio can increase the marriage quality, which decreases divorce rates; on the other hand, a higher sex ratio can increase outside options, which destabilizes marriage. The model demonstrates the impacts of the two forces, illustrating that it is possible to derive either positive or negative outcomes from the same theoretical framework under varying conditions. The model is related to previous papers such as Burdett and Mortensen (1998) where workers search both on and off the job.

We assume that all women are identical, and we consider the decisions of an infinitely lived woman in continuous time. At a moment in time, each woman is either single (state 0) or married (state 1). At random time intervals, a woman meets a new potential partner. The arrival rates depend on a woman's current state, i.e., single or married. Let λ_i, $i \in \{0, 1\}$ represent the parameter of a Poisson arrival process, which denotes the arrival rate of new partners while a woman is currently in state i.

A higher sex ratio means higher arrival rates of new potential partners for both single and married women. Namely, λ_0 and λ_1 increase with the sex ratio. In the following analyses, we first derive comparative statics with respect to λ_0 and λ_1. Then, we discuss the implications with different relative impacts of the sex ratio on λ_0 and λ_1.

Women are assumed to search among potential partners randomly. The quality of a new partner is assumed to be the realization of a random draw from a distribution F if a woman is single and a distribution G if a woman is married. Women must respond to new partners as soon as they arrive; there is no recall.

Given the framework briefly outlined above, the expected discounted lifetime utility when a woman is single, U, can be expressed as the solution to the following equation:

$$rU = b + \lambda_0 \int \max\{V(w) - U, 0\} dF(w)$$

In other words, the opportunity cost of searching while single, the interest on its asset value, is equal to utility while single plus the expected capital gain attributable to searching for an acceptable partner where acceptance occurs only if the value of marriage, $V(w)$, exceeds that of continued search.

Similarly, the expected lifetime utility of a married woman currently with a partner of quality w solves the following:

$$rV(w) = w + \lambda_1 \int \max\{X(\tilde{w} - c) - V(w), 0\} dG(\tilde{w})$$

where $X(\tilde{w}-c)=\frac{\tilde{w}-c}{r}$. We assume that people divorce at most once, potentially due to the substantial divorce cost.[37]

We solve the model backward. For married women with a husband of quality w, there exists a reservation quality, w_{R1}, such that $X(w_{R1}-c)=V(w)$. w_{R1} satisfies the following:

$$w_{R1}=c+w+\frac{\lambda_1}{r}\int_{w_{R1}}(\tilde{w}-w_{R1})dG(\tilde{w}) \tag{5}$$

Equation 5 is the implicit function expression of w_{R1}. w_{R1} is thus a function of w and λ_1. c and r also affect w_{R1}, and we assume that they are exogenous and take them as given. Denote w_{R1} as $w_{R1}(w,\lambda_1)$. It can be derived that:

$$\frac{\partial w_{R1}(w,\lambda_1)}{\partial w}=\frac{1}{1+\frac{\lambda_1}{r}(1-G(w_{R1}))}>0$$

$$\frac{\partial w_{R1}(w,\lambda_1)}{\partial \lambda_1}=\frac{\frac{1}{r}\int_{w_{R1}}(\tilde{w}-w_{R1})dG(\tilde{w})}{1+\frac{\lambda_1}{r}(1-G(w_{R1}))}>0$$

For single women, there exists a reservation quality, w_{R0}, such that $V(w_{R0})=U$. w_{R0} satisfies the following:

$$rU=rV(w_{R0})=w_{R1}(w_{R0},\lambda_1)-c=b+\frac{\lambda_0}{r}\int_{w_{R0}}(w_{R1}(w,\lambda_1)-w_{R1}(w_{R0},\lambda_1))dF(w)$$

Rearrange the equation, we get the implicit function that defines the value of w_{R0}:

$$w_{R1}(w_{R0},\lambda_1)=c+b+\frac{\lambda_0}{r}\int_{w_{R0}}(w_{R1}(w,\lambda_1)-w_{R1}(w_{R0},\lambda_1))dF(w) \tag{6}$$

It can be derived that:

$$\frac{\partial w_{R0}}{\partial \lambda_0}=\frac{\frac{1}{r}\int_{w_{R0}}(w_{R1}(w,\lambda_1)-w_{R1}(w_{R0},\lambda_1))dF(w)}{\frac{\partial w_{R1}(w_{R0},\lambda_1)}{\partial w_{R0}}\left(1+\frac{\lambda_0}{r}(1-F(w_{R0}))\right)}>0$$

which means a higher arrival rate in the first-marriage market increases the reservation value.

[37] This assumption is consistent with data. According to CFPS 2010 data, less than 1% of people ever married more than twice.

Additionally, it can be derived that:

$$\frac{\partial w_{R0}}{\partial \lambda_1} = -\frac{\frac{\partial w_{R1}(w_{R0},\lambda_1)}{\partial \lambda_1} - \frac{\lambda_0}{r}\int_{w_{R0}}\left(\frac{\partial w_{R1}(w,\lambda_1)}{\partial \lambda_1} - \frac{\partial w_{R1}(w_{R0},\lambda_1)}{\partial \lambda_1}\right) dF(w)}{\frac{\partial w_{R1}(w_{R0},\lambda_1)}{\partial w_{R0}}\left(1+\frac{\lambda_0}{r}(1-F(w_{R0}))\right)} < 0 \tag{7}$$

We then analyze the duration of the first marriage, which is related to divorce. Specifically, a longer duration of the first marriage corresponds to a lower divorce rate. The average duration of the first marriage has the following expression:

$$E(\text{duration}) = \int_{w_{R0}} \frac{1}{\lambda_1(1-G(w_{R1}(w,\lambda_1)))} dF(w)/[1-F(w_{R0})]$$

It can be derived that:

$$\begin{aligned}\frac{\partial E(\text{duration})}{\partial \lambda_0} =& \frac{\partial w_{R0}}{\partial \lambda_0}\frac{f(w_{R0})}{(1-F(w_{R0}))^2}\frac{1}{\lambda_1}\times \\ & \int_{w_{R0}}\left(\frac{1}{1-G(w_{R1}(w,\lambda_1))} - \frac{1}{1-G(w_{R1}(w_{R0},\lambda_1))}\right) dF(w)\end{aligned} \tag{8}$$

It can be shown that $\frac{\partial E(\text{duration})}{\partial \lambda_0} > 0$.

$$\begin{aligned}\frac{\partial E(\text{duration})}{\partial \lambda_1} =& \int_{w_{R0}} \frac{\partial \frac{1}{\lambda_1(1-G(w_{R1}(w,\lambda_1)))}}{\partial \lambda_1} dF(w)\frac{1}{1-F(w_{R0})} \\ &+ \frac{\partial w_{R0}}{\partial \lambda_1}\frac{f(w_{R0})}{(1-F(w_{R0}))^2}\frac{1}{\lambda_1}\times \\ & \int_{w_{R0}}\left(\frac{1}{1-G(w_{R1}(w,\lambda_1))} - \frac{1}{1-G(w_{R1}(w_{R0},\lambda_1))}\right) dF(w)\end{aligned} \tag{9}$$

It can be shown that $\frac{\partial E(\text{duration})}{\partial \lambda_1} < 0$ if the distribution $G(\tilde{w})$ is log-concave.

The equation suggests that there are two forces through which λ_1 affects the duration of the first marriage. Firstly, λ_1 is the arrival rate and affects the duration of the first marriage directly. Secondly, expecting a high arrival rate after marriage, women lower their reservation value for the first marriage, which affects the duration. In other words, λ_1 affects the duration indirectly by affecting the reservation value for the first marriage, w_{R0}.

Denote the sex ratio as s, $\eta_0 = \frac{\partial \lambda_0}{\partial s}$ and $\eta_1 = \frac{\partial \lambda_1}{\partial s}$, we can derive the following:

$$\frac{\partial w_{R0}}{\partial s} = \frac{\partial w_{R0}}{\partial \lambda_0}\eta_0 + \frac{\partial w_{R0}}{\partial \lambda_1}\eta_1$$

Therefore, we have $\frac{\partial w_{R0}}{\partial s} > 0$ as long as $\frac{\eta_0}{\eta_1} > -\frac{\frac{\partial w_{R0}}{\partial \lambda_1}}{\frac{\partial w_{R0}}{\partial \lambda_0}}$. Given that most people are in their first marriages and the observed average quality of marriages is the average of qualities above w_{R0}, namely $E[w|w > w_{R0}]$, $\frac{\partial w_{R0}}{\partial s} > 0$ means the overall marriage quality increases with a higher sex ratio.

Additionally, Equation (8) suggests that $\frac{\partial E(\text{duration})}{\partial \lambda_0}$ can be expressed as $A\frac{\partial w_{R0}}{\partial \lambda_0}$ where A is a positive number. Equation (9) suggests that $\frac{\partial E(\text{duration})}{\partial \lambda_1}$ can be expressed as $(A\frac{\partial w_{R0}}{\partial \lambda_1} - B)$ where B is also a positive number. Combined together, we can derive the following:

$$\frac{\partial E(\text{duration})}{\partial s} = \frac{\partial E(\text{duration})}{\partial \lambda_0}\eta_0 + \frac{\partial E(\text{duration})}{\partial \lambda_1}\eta_1 = A\left(\frac{\partial w_{R0}}{\partial \lambda_0}\eta_0 + \frac{\partial w_{R0}}{\partial \lambda_1}\eta_1\right) - B\eta_1$$

Therefore, we have $\frac{\partial E(\text{duration})}{\partial s} < 0$ as long as $\frac{\eta_0}{\eta_1} < \frac{B/A - \frac{\partial w_{R0}}{\partial \lambda_1}}{\frac{\partial w_{R0}}{\partial \lambda_0}}$. $\frac{\partial E(\text{duration})}{\partial s} < 0$ means the duration of first marriages decreases when the sex ratio increases, and thus divorce rates increase with a higher sex ratio.

Given that $\frac{\partial w_{R0}}{\partial \lambda_1} < 0$, there exist values of η_0 and η_1 that satisfy both $\frac{\eta_0}{\eta_1} > -\frac{\frac{\partial w_{R0}}{\partial \lambda_1}}{\frac{\partial w_{R0}}{\partial \lambda_0}}$ and $\frac{\eta_0}{\eta_1} < \frac{B/A - \frac{\partial w_{R0}}{\partial \lambda_1}}{\frac{\partial w_{R0}}{\partial \lambda_0}}$.

The framework presented in this section implies that, with certain parameters, we can observe both: 1) the marriage quality increases with the sex ratio, and 2) the divorce rates increase with the sex ratio.

B Additional Tables and Figures

B.1 Descriptive Statistics

Table B.1: Descriptive Statistics

Variable	Mean	Std. Dev.	Min	Max
Panel A: Province-level Data				
(# of divorce cases / population aged 15-64)×1000‰	2.503	2.118	.001	9.479
(# of divorce cases / # of marriage cases)×100	28.449	14.96	6.872	80.812
Sex ratio	1.056	.051	.946	1.224
ln(GDP per capita)	9.947	.937	7.719	12.009
Average household size	3.258	.485	2.222	6.79
Average year of education	8.309	1.375	2.948	12.782
Panel B: Census Data				
Divorce	.016	.127	0	1
Sex ratio	1.031	.077	.826	1.259
Han Ethnicity	.913	.282	0	1
Employed	.775	.417	0	1
Migrant	.269	.444	0	1
Urban	.353	.478	0	1
Having a son	.651	.477	0	1
Number of children	1.297	.762	0	9
Panel C: Individual Survey Data				
Divorce	.028	.166	0	1
Sex ratio	1.039	.065	.842	1.325
Han Ethnicity	.912	.283	0	1
Employed	.753	.431	0	1
Migrant	.093	.291	0	1
Urban	.472	.499	0	1
Cohabit	.122	.327	0	1
Having a son	.692	.462	0	1
Number of children	1.581	.851	0.000	7
Number of male sibling(s)	1.485	1.181	0.000	7
Number of female sibling(s)	1.364	1.302	0.000	8
Income	8348.388	17044.929	0.000	800000

NOTE: The province-level data has 744 observations except for the divorce-marriage ratio, which has 651 observations. The census data has 1,150,344 observations; the CFPS survey data has 8,717 observations.

B.2 Robustness Tests

Table B.2: Robustness Tests

	Age 25-45	[-6,+1]	[-8,+2]/[-7,+3]	Wave-specific sex ratio
	(1)	(2)	(3)	(4)
Sex Ratio	0.022***	0.016***	0.020***	0.021***
	(0.006)	(0.004)	(0.004)	(0.005)
Obs.	1,387,342	1,150,344	1,150,344	1,231,916
Adj. R-sq	0.020	0.018	0.018	0.018
Mean of Dep. Var.	0.019	0.016	0.016	0.016

NOTE: This table shows the results of the first set of robustness tests. Column (1) reports the estimations using the sample of women aged 25-45. Columns (2) and (3) report results using [-6, +1] and [-8,+2]/[-7,+3] age gaps for the calculation of the sex ratio. Column (4) shows the results using census wave-specific sex ratio. Specifically, we construct the sex ratio by aggregating data from both the present and preceding waves of censuses. For example, to construct the sex ratio for observations in the 2010 census, we calculate the sex ratio pooling 2000, 2005, and 2010 waves. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

Table B.3: Robustness Tests (Cont.)

	Industry & occupation	Province × year FE	Fake divorce	Exclude migrants	County-level sex ratio
	(1)	(2)	(3)	(4)	(5)
Sex Ratio	0.018***	0.016***	0.019***	0.017***	0.013***
	(0.005)	(0.004)	(0.004)	(0.005)	(0.003)
Obs.	891,729	1,150,346	1,150,346	840,431	632,769
Adj. R-sq	0.021	0.018	0.018	0.020	0.018
Mean of Dep. Var.	0.016	0.016	0.016	0.017	0.019

NOTE: This table shows the results of the second set of robustness tests. Column (1) reports the estimations controlling for census wave-specific occupation and industry fixed effects. Column (2) reports the results controlling for province × year fixed effects. Column (3) reports results controlling for the implementation of fake divorce. Column (4) shows the results excluding migrants. Column (5) shows the results using an alternative sex ratio measure. Specifically, we restrict the sample to the census years 2010 and 2015 because the year 2005 does not have information on the residential county. For 2010 and 2015, we replaced the prefecture-level sex ratio with the county-level sex ratio if the women lived in rural areas. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

B.3 Results using Instrumental Variables

We have experimented with three types of instruments. First, previous literature has documented that the land reform in China in the 1980s has increased the male-to-female ratio (Almond et al., 2019). We use the indicator for whether the land reform has begun in the women's living prefecture five years before her birth year as the instrument for the sex ratio in her birth cohort.

Secondly, the spread of ultrasound machines facilitated sex selection (see, e.g. Hesketh and Xing, 2006). We use the indicator for the presence of ultrasound machines in the capital of each woman's province of residence in her birth year as an instrument for the sex ratio.[38]

Finally, as migrants remain primarily men in our sample period, increases in migration can decrease the origin's sex ratio. Correspondingly, we use the FDI in the main flow-in prefectures in China: Beijing, Shanghai, and Shenzhen, weighted by the inverse distance to the women's prefecture of residence as an instrument for sex ratio. Specifically, for women living in prefecture p born in year y, we construct the following instrument:

$$IV_{py} = \sum_{c} \frac{lnFDI_{c,y+20}}{distance_{pc}}$$

where c can be Beijing, Shanghai, or Shenzhen, $lnFDI_{c,y+20}$ is the log of FDI in prefecture c in year $y+20$, $distance_{pc}$ is the distance in miles between prefecture p and c. The sample excludes women living in Beijing, Shanghai, and Shenzhen. The first stage is expected to be negative as migration decreases the origin's sex ratio.

Table B.4 presents the results using all three instruments. The IVs have relatively weak first stages. Only lnFDI/distance has an F-stat larger than 10, the rule of thumb. While the second-stage estimations are noisy, all the confidence intervals cover the OLS estimation. We have experimented with various variations for all these instruments, including age thresholds, distance measures, etc. The results remain quantitatively similar.

[38]The data on the presence of ultrasound machines comes from Almond et al. (2019).

Table B.4: IV Results

	Sex ratio	Divorce	Sex ratio	Divorce	Sex ratio	Divorce
	(1)	(2)	(3)	(4)	(5)	(6)
Sex ratio		0.119**		0.012		0.044**
		(0.055)		(0.046)		(0.020)
Land Reform	0.027***					
	(0.008)					
Ultrasound			0.014**			
			(0.006)			
lnFDI/distance					-5.499***	
					(1.339)	
F-stat	11.04		5.467		16.865	
Obs.	1,071,983	1,150,344	1,048,586	1,048,586	892,060	892,060

NOTE: Odd columns are the first stages, and even columns are the second stages for IV estimations. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

B.4 Gender Imbalance and Remarriage of Men

Table B.5: Gender Imbalance and Remarriage of Men

1{Ever-remarry}	Divorced	Divorced & Widowed
	(1)	(2)
Sex ratio	-1.186	-0.909
	(0.799)	(0.737)
Obs.	366	483
Adj. R-sq	0.224	0.194
Mean of Dep. Var.	0.488	0.447

NOTE: Prefecture, birth year, education, number of children, male and female siblings fixed effects are controlled for all columns. All the control variables for the census regressions and the indicator for cohabitation before marriage are controlled for all columns. Standard errors clustered at the prefecture level are reported in parentheses. ***p<0.01, **p<0.05, *p<0.1.

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