Premarital Cohabitation and Marital Disruption: Using Propensity Score Matching to Analyze the Cohabitation Effect in Urban China

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Abstract

This article uses the propensity score matching (PSM) method to analyze the cohabitation effect in urban China: the relationship between premarital cohabitation and marital disruption. Using data on urban individuals married during 1978-1999 and 2000-2010 from Chinese Family Panel Studies in 2010, this article matches married ones with cohabitation histories prior to first marriage—whether cohabited with the spouse or not, and applies Cox Proportional Hazards Models to explore how premarital cohabitation affects an individual's subsequent divorce hazards of first marriage. This research demonstrates that both causal and diffusive theoretical frameworks on the cohabitation effect initiated by western scholars apply to the context of contemporary China. Specifically, the cohabitation effect existed when cohabitation was uncommon during 1978-1999 which indicates that cohabitation increased individuals' divorce hazards. This effect, however, disappeared when cohabitation became popular during 2000-2010. Besides, those selective factors such as education, income, household registration (*hukou*) and childhood experience also could partly explain the cohabitation effect when cohabitation was uncommon. An imputation-based sensitive analysis shows that this article's conclusion is robust, though the existence of unobserved covariates confounds the relationship between premarital cohabitation and marital disruption.

Key words: premarital cohabitation, marital disruption, propensity score, optimal matching

Introduction

For the past several decades, cohabitation—a living arrangement of heterosexual unmarried couples—has become very prevalent in western countries; the majority of marriages and remarriages begin with cohabitations. In the United States, the number of cohabiting couples grew from 500,000 in 1970 to 7.8 million in 2012 (U.S. Census Bureau 2012) and the percentage of marriages preceded by cohabitation for those married between 1965 and 1974 rose from 10% to more than 50% for those married between 1990 and 1994 (Bumpass & Lu 2000; Bumpass & Sweet 1989). Major empirical studies by demographers and family sociologists have also shown that premarital cohabitation is strongly associated with a greater risk of marital disruption in western contexts, which is defined as the cohabitation effect (Axinn & Thornton 1992; Bumpass, Sweet, & Cherlin 1991; Cohan & Kleinbaum 2002; Lu et al. 2012; Stanley, Rhoades, & Markman 2006; Teachman & Polonko 1990).

This social phenomenon, however, is receiving little attention by Chinese scholars despite the prevalence of premarital cohabitation in contemporary China. This article attempts to explore whether the cohabitation effect exists in the Chinese context. By applying the PSM method and Cox Proportional Hazards Models, this study aims to determine if there is a causal relationship between premarital cohabitation and subsequent divorce hazards among urban married couples.

Premarital Cohabitation and Marital Disruption

Meaning of Cohabitation

The meaning of cohabitation is critical to understanding the relationship between cohabitation and marriage. Early research on cohabitation and marriage was mainly framed in two views, cohabitation as a prelude to marriage and cohabitation as an alternative to marriage (Xie et al. 2003). Lately these two views have expanded into six ideal types of cohabitation which have different impacts on marriage or union formation: (a) marginal, (b) prelude to marriage, (c) stage in the marriage process, (d) alternative to single, (e) alternative to marriage, and (f) indistinguishable from marriage (see Heuveline & Timberlake 2004). In fact, people in the United States currently regard cohabitation as either a testing ground for marriage or an alternative to marriage (Smock 2004). From the trial marriage perspective, cohabitation actually plays a protective role in the subsequent marriage because it excludes those divorce-prone or unmatched potential spouses. In contrast, individuals who view cohabitation as an alternative to marriage are less likely to get married or just move from one cohabiting relationship to another (Casper & Bianchi 2002). For an individual who sees cohabitation as a stage in the marriage process, they may be more possible to get married for raising children within marriage especially when institutional arrangements or cultural sanctions against out-of-wedlock childbearing exist (Heuveline & Timberlake 2004). A better understanding of the evolution of cohabitation will be beneficial in capturing the nature of cohabitation in contemporary China.

Selective, Causal or Diffusive?

Prior studies have revealed that cohabitation is strongly associated with marital disruption (Cohan & Kleinbaum 2002; Kamp Dush, Cohan, & Amato 2003; Stanley, Rhoades, & Markman 2006). Three approaches have arisen to explain the links between premarital cohabitation and marital disruption.

The selective approach contends that cohabitors are a select group of individuals who differ in salient ways from non-cohabitors. More specifically, these pre-cohabiting selective characteristics such as younger age, less religiosity, parental divorce, lower socio-economic status, or liberal attitudes toward premarital sex make cohabitors less committed to marriage and more approving of divorce (Axinn & Thornton 1992; Booth & Johnson 1988; Lillard, Brien, & Waite 1995; Thomson & Colella 1992; Woods & Emery 2002). Thus, the link between cohabitation and marital disruption is explained by the types of cohabitors with some divorce-prone demographic characteristics.

On the contrary, the causal approach assumes that cohabitation itself significantly increases the risk of marital disruption beyond those selective features which bring them into cohabitation (Kamp Dush et al. 2004). According to this perspective, cohabitation changes individuals' commitment to marital relationships in ways that undermine subsequent marital stability, thus leading to greater acceptance of divorce (Stafford, Kline, & Rankin 2004; Stanley, Whitton, & Markman 2004). Early research by Kamp Dush, Cohan, and Amato (2003) using data of two U.S. marriage cohorts (1964-1980 & 1981-1997) from Marital Instability Over the Life Course has shown that cohabitors in both cohorts continue to exhibit worse marital quality and greater marital disruption after selection factors for cohabitation and subsequent marital disruption were controlled.

Contrary to the two approaches discussed above, the diffusive approach argues that the relationship between cohabitation and marital disruption changes across different cohorts. It posits that the cohabitation effect decreases or even disappears as cohabitation becomes more prevalent (De Vaus, Qu, & Weston 2005; Hewitt & De Vaus 2009; Manning & Cohen 2012). Currently, married couples preceded by cohabitation have become less selective and cohabitation gradually becomes a typical union formation (Reinhold 2010). Previous research by Reinhold (2010), which used data from National Survey of Family Growth (2002), has corroborated that no cohabitation effect exists among those female marriage cohorts married after 1993. Likewise, the NFSG2006-2008 data has also presented the same result among those married after 1996 when 61% of married individuals have cohabitation experiences that there is no association between cohabitation and marital disruption (Manning & Cohen 2012).

Nonetheless, because of the technical constraints of causal inferences, it is not clear whether the association between cohabitation and marital disruption is causal or selective. Recently, some scholars have attempted to use the PSM method to examine the selective, causal and diffusive perspectives of the cohabitation effect in western contexts. For instance, Lu et al. (2012) used data from National Survey of Families and Households with the propensity score optimal matching method to substantiate that the selection effect played a large role in 1987–1988 when cohabitation was uncommon but disappeared in 2001–2003 when cohabitation became prevalent, and that the causal effect of cohabitation on marital disruption was strong among serial cohabitors but weak among one-time cohabitors with the spouse only. Therefore, the result indicates that both selection and causation help explain the cohabitation effect.

However, it is still unclear whether this cohabitation effect exists in non-western contexts. Thus, this article attempts to extend Lu et al.'s results into the Chinese context and investigate the cohabitation effect in urban China by using data from Chinese Family Panel Studies in 2010.

Research Data and Methods

Data

The data are from Chinese Family Panel Studies (CFPS), funded by 985 Program of Peking University and carried out by the Institute of Social Science Survey of Peking University in 2010. CFPS2010 is a national probability sample of 33,600 adult respondents from rural and urban China including 8,481 urban respondents with marriage experiences and 869 urban respondents with cohabitation experiences prior to first marriage. Detailed first marriage and cohabitation histories (whether respondents cohabited with the spouse prior to the first marriage or not) are collected from urban family adults (hereinafter cohabitation is generally referred to cohabiting with the spouse prior to the first marriage). Since cohabitation was very rare during the Maoist' Period (1949-1976), this article only focuses on urban individuals with marriage experience after the Economic Reform Period (after 1978). The total samples are divided into marriage cohort 1978-1999 when cohabitation was rare and marriage cohort 2000-2010 when cohabitation was prevalent. Furthermore, to obtain the net cohabitation effect, individuals with cohabitation experiences prior to first marriage are in one treatment group and those without cohabitation experiences are in another.

Propensity Score Analysis

The propensity score, the conditional probability of assignment to a particular treatment given a vector of observed covariates, is used to balance the selective biases of observed covariates between a treatment group and a control group if the treatment is binary (Rosenbaum & Rubin 1983). This method has been widely used in observational studies among various fields such as sociology, social policy and public health (e.g., Harding 2003; Lu et al. 2012; Morgan & Winship 2007).

In a binary treatment effect design, Y^T (T=1if treated; 0, otherwise) denotes the potential outcomes for a subject under binary treatment conditions and vector **X** denotes a set of observed covariates. Thus, a valid causal inference relies on the strongly ignorable treatment assignment assumption: $T \perp Y^T \mid e(\mathbf{X})$, where e(**X**) is the balancing score defined as $P(T=1|\mathbf{X})$. This assumption assures that individual i can be randomly assigned into a treatment group, which leads to a random-like situation (Rosenbaum & Rubin 1983). In this article, whether a person chooses to cohabit before first marriage can be viewed as a non-random experiment because the differences in preexisting characteristics, such as education, parental backgrounds, and socio-economic status would affect the treatment effect on marital disruption (Lu et al. 2012). Hence, it is important to remove those observed biases that influence the probability of cohabitation to estimate the net cohabitation effect on marital disruption.

This article employs a binary logistic model to estimate the probability of accepting treatment(cohabiting with the spouse prior to first marriage)：

$$\log (P(T=1) / (1-P(T=1)))=\alpha+\beta x_i+\varepsilon$$

The suggestion given by Rosenbaum & Rubin (1985) to use the logistic transformation of predicted probability log ((1- e (**X**))/e (**X**)) as propensity score is accepted by this article due to its proximity to normal distribution.

Optimal Matching

Matching is "a method of sampling from a large reservoir of potential controls to produce a control group of modest size in which the distribution of covariates is similar to the distribution in the treated group" (Rosenbaum & Rubin 1983). After getting the propensity score, this scalar balancing score can be used to balance distributions of covariates by matching cohabitors and non-cobabitors among urban married individuals.

This article takes the optimal pair matching as its matching strategy implemented by *optmatch* package in *R* developed by Hansen through the use of network flow theory (Hansen & Klopfer 2006). Put simply, optimal matching is a process to develop S matched sets consisting of treated (T_s) and control (C_s) subjects that minimizes the total distance Δ of propensity scores defined as

$$\Delta = \sum_{s=1}^{S} \omega(|T_s|,|C_s|)\delta(T_s, C_s)$$

Where $|C_s|$ denotes the number of control subjects and $|T_s|$ the number of treated participants in the S^{th} stratum. $\omega(|T_s|,|C_s|)$ is a weight function of the sizes of the treated and control subjects and $\delta(T_s, C_s)$ is the average distance of treated and control participants within the S^{th} stratum (Rosenbaum 2002:308-310). In an optimal pair 1: k matching, each treated participant matches to k controls (Guo & Fraser 2009:151). Considering sample sizes of two marriage cohorts, optimal pair 1:4 matching for cohort 1978-1999 and optimal pair 1:1 matching for cohort 2000-2010 are adopted in this article.

Event History Analysis

After obtaining the matched data, this article applies Cox Proportional Hazards Models to estimate hazard ratios of divorce on first marriage. The final analytic samples of this paper consist of ever-married urban individuals who either were in their first marriages in 2010 or had experienced the dissolution of their first marriages. The duration (years) of the first marriage and whether it ended up with divorce at the time of survey are checked to get the dependent variable which is “if the first marriage broke up at time t while it still remained stable at t-1.” The PSM technique is applied to remove the selection bias of cohabitation from those observed covariates such as education, age, gender, party membership, income, family background, and childhood experience, followed by the application of Cox models using the matched data. All demographic preexisting features as well as respondents’ first marriage ages are added into Cox models as control variables. Furthermore, because parental backgrounds such as education and marital status are missing in the data, this study also conducted the sensitive analysis to examine the confounding effects of those unobserved covariates.

Research Results

Change in Prevalence of Cohabitation and Divorce in Contemporary China

Table 1 (see appendix) shows the distribution for cohabitation in different marital periods. For the urban sample, the percentage of respondents who cohabited with their spouses prior to first marriage increased from 1.82% for those who married before 1978 to 7.56% for those who married between 1978 and 1999, and finally, to 28.36% for those who married between 2000 and 2010. The national sample is generally consistent with this pattern. Obviously, cohabitation has gradually become prevalent after the Late Reform Period due to economic development and changing attitudes toward marriage and premarital sex.

Table 2 (see appendix) presents the distribution for the duration of cohabitation. For the national sample, the duration of cohabitation in three marital periods all displays that the majority of cohabitors chose to get married in less than 12 months and only a few couples cohabited more than 1 year. This may show that cohabitation is primarily considered as a precursor of marriage or as one marital stage rather than an alternative to marriage. However, for the urban sample, the percentage of respondents who cohabited with the spouses more than 1 year increased from 12.50% for those married before 1978 to 16.23% for those married in 1978-1999, and to 20.49% for those married in 2000-2010. This may show that more and more urban individuals prefer a long cohabiting relationship with their partners to an immediate marriage relationship.

With respect to divorce of the first marriage, for the national sample, the percentage of divorcers among those married after 1978 presented in table 3 (see appendix) was greater than that of those married before 1978, which implies that general social acceptance of divorce in China has also increased for the last several decades. For urban individuals, with higher divorce rates of first marriage, they present a more divorce-prone characteristic, which indicates that urban citizens were becoming more approving of divorce since the Economic Reform.

Propensity Score Matching Analysis[1]

[1] Propensity score matching analysis mainly involves imbalance checking of covariates before matching, propensity score estimation, identification of matched sets with matching algorithm, and balance checking of covariates after matching.

The results of descriptive analysis of the covariates on cohabitation experience are presented in table 4 (see appendix) which demonstrate substantial imbalance between cohabitors and non-cohabitors. This article mainly discusses the results summarized in table 4 on urban marriage cohort 1978-1999. The absolute standardized differences[2] (a measure of imbalance; see Haviland 2007, 2008) before matching (*dx*) show that covariates such as education, income, birth year, party membership, living area, and *hukou* when respondents were 12 years old are not balanced in both marriage cohorts. For instance, the largest absolute standardized difference of birth year (1980-1987) is up to 0.602 for marriage cohort 1978-1999 which indicates a large imbalance.

The logistic regression is used to predict the propensity score which is the individual's possibility of belonging to the treatment group rather than the control group. Taking urban marriage cohort 1978-1999 as an example, the logistic regression shows that during the Early Economic Reform Period when cohabitation was uncommon, CCP members are less likely to cohabit before first marriage, probably because of political or moral constraints, while people who possess economic advantages are much prone to cohabit with their partners because they are more capable to afford the cost of living together. The regression also indicates that individuals living in eastern China are more likely to cohabit, probably due to a more mature economy or changing attitudes. Those born before 1960 have more conservative attitudes toward cohabitation with a prude virginity ideology during the Maoist' Period (Li 1991). Marriage cohort 2000-2010 generally follows the same pattern.

Optimal pair matching is implemented to balance selective biases from covariates between cohabitors and non-cohabitors for obtaining our final analytical samples. Figure 1 (see appendix) presents the distribution for propensity scores before matching. It shows that, in spite of deleting some cases, the overlaps of propensity scores are sufficient enough to get our matched samples for two marriage cohorts.

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[2] Haviland et al. (2007, 2008) developed the imbalance checking method of absolute standardized difference in covariate means: *dx* for use before matching and *dxm* for use after matching. *dx* and *dxm* can be explained as the difference between treated and control groups on X in terms of the standard deviation unit of X. Scholars expect to see *dx* > *dxm* because the sample balance should be improved after matching (Guo & Fraser 2009:157-158).

The balance checking through absolute standardized differences after matching (*dxm*) presented in table 4 shows that optimal pair matching balances those distributions of covariates very well, and then this article uses matched data to run Cox Proportional Hazards Models with all those covariates discussed above including age of first marriage to explore the cohabitation effect.

The Analysis of Cox Proportional Hazards Models[3]

Table 5 (see appendix) shows the hazard ratios of marital disruption of different marriage cohorts before matching. It reveals that the cohabitation effect in urban samples changes across different marriage cohorts. For those married between 1978 and 1999, prior cohabitation experiences increased the divorce hazards of the subsequent first marriage significantly by 92%. Not surprisingly, the cohabitation effect disappeared when cohabitation became prevalent during the period of 2000-2010. Though not statistically significant, cohabitation even reduced the subsequent divorce hazard by 28% compared with non-cohabitors. According to Cox models before matching, it can be summarized that when cohabitation was not popular, cohabitation indeed increased its subsequent risk of marital disruption but this effect decreased or even disappeared when cohabitation became commonplace. However, it is still unreasonable to conclude that there is a causal relationship between cohabitation and marital disruption because of those confounding selective biases just as is presented in table 4.

Hazard ratios of the matched data are also predicted and presented in table 6 (see appendix). The results are consistent with what are discussed with original samples. After optimal pair matching, the divorce hazard ratio of first marriage, 1.60, for those marred between 1978 and 1999, were slightly smaller than that before matching, which indicates that cohabitation increased the marital risk by 60% compared with non-cohabitors. This might also show that selective factors partly account for subsequent marital disruption when

[3] This article only focuses on the cohabitation effect on urban individuals. Actually, the author has also conducted a detailed analysis of this effect on rural individuals. The rural is basically consistent with the urban.

cohabitation was uncommon. The cox models after optimal pair matching indicated that the cohabitation effect disappeared or even reduced divorce hazards. For instance, the divorce hazard of cohabitors from urban areas who married between 2000 and 2010 was 23% lower than that of non-cohabitors.

According to those discussions above, it can be concluded that premarital cohabitation is causally but conditionally associated with divorce hazards of first marriage. Cohabitation did increase urban individuals' divorce hazards during the Early Economic Reform Period (1978-1999) when cohabitation was uncommon, but the cohabitation effect disappeared during the Late Economic Reform Period (2000-2010) when cohabitation became relatively prevalent. This result shows that the diffusive perspective is also well applied in contemporary China. More importantly, the cohabitation experience even reduced the urban individuals' divorce hazards though not statistically significant. One possible explanation is that cohabitation in contemporary China has become a test ground of marriage which excludes those unsuitable partners, thus leading to a more stable marital relationship among those who withstand the trials.

Sensitive Analysis

Propensity score analysis, though efficiently balancing observed covariates, is not able to control unobserved biases. This article's models include a series of covariates that may affect premarital cohabitation and divorce hazards of first marriage. Just as mentioned above, a valid causality is based on the fact that those covariates could efficiently explain individuals' differences of cohabiting experiences. However, if those unobserved covariates bias the assignment of treatment, the causality will not be sound. Therefore, the sensitive analysis must be carried out to examine whether the assumption is violated.

Sensitive analysis assumes that an unobserved covariate, *U*, may potentially change the odds of receiving treatment to certain degree, Γ, given that all other covariates are observed. The sensitivity analysis examines how much hidden bias can be present—that is, "how large must Γ be—before the qualitative conclusion begins to change" (Rosenbaum 2005: 1809). This article only focuses on how large Γ could make the cohabitation effect change.

This article adopts the Lu et al.'s method of conducting sensitivity analysis for event history analysis which is to run an imputation-based sensitivity analysis by treating *U* as an unobserved covariate, imputing *U* based on a hypothetical model for each case, and including *U* as a covariate in the final Cox models (Lu et al. 2012).

In this article, *T* (1or 0) denotes individual' cohabiting experiences with the spouse prior to first marriage, *Y* indicates individual's survival time of first marriage (to simplify, if the survival outcome is either less than or equal to median survival time to divorce, *Y*=1; 0, otherwise), and censoring variable *C* shows whether the first marriage ends in divorce and assumes that *C* and *U* are independent. The following equation, a logistic model, is implemented to generate the unobserved covariate *U*:

$$\text{logit}\ (P)=\beta_1 T+\beta_2 Y\cdot C$$

Where β_1 denotes the log odds ratio of *U* being in cohabiting with the spouse prior to first marriage group (*T* = 1) rather than in non-cohabitation group (*T* = 0). β_2 is used to assess the relationship between U and the timing of marital disruption, which is also the log odds ratio of U being in the shorter marriage survival group (*Y*=1) rather than in the other (*Y*=0). Thus, this article could use β_1 and β_2 together to evaluate potential deviation from the strong ignorable assignment assumption. For instance, $(e^{\beta 1}, e^{\beta 2})$ = (1, 1) implies that unobserved covariate *U* is independent with premarital cohabitation and marital disruption, and if (2, 2), *U* is associated with premarital cohabitation (two times as likely in odds) and shorter marriage survival time (two times as likely in odds). In fact, *U* as an unobserved covariate can be treated as a typical missing data problem. Hence, the multiple imputation technique can be implemented to generate *U*, allowing $e^{\beta 1} and\ e^{\beta 2}$ varying between 1 and 10 (1, no association with *U*; 10, a very strong association with *U*). According to Lu's suggestion (2012), each combination of $e^{\beta 1}$and $e^{\beta 2}$ is performed 10 times to avoid underestimating the variance. Besides, this paper merely conducts sensitive analysis for marriage cohort 1978-1999 because the cohabitation effect of marriage cohort 2000-2010 is not significant.

Sensitive analysis concentrates on to what certain degree can research conclusion change qualitatively, thus this article applies commonly used 5 percent (p=.05) as the p value for a qualitative change of the research conclusion. Figure 2 (see appendix) shows that, only with very large odds ratios, say, $(e^{\beta 1}, e^{\beta 2})$ = (3, 3), can this article's conclusion be qualitatively changed. In other words, sensitive analysis of optimal pair matching indicates that the qualitative conclusion could be altered only if hidden bias U is strongly associated with premarital cohabitation and shorter survival time of first marriage. As a result, it can be concluded that this article's analysis is robust. If more covariates such as parents' information can be balanced, this article's conclusion will be more robust. Of course, this is not to deny that those hidden biases especially from parental backgrounds and childhood experiences still might significantly affect Chinese individuals' cohabiting choices which are worth further study.

Conclusion and Discussions

This article's research findings indicate that for the last several decades cohabitation has become more and more commonplace in both Urban and Rural China. Why is cohabitation prevailing in contemporary China compared with previous periods? One possible explanation can be related to the individualization of Chinese society (Yan 2009). In traditional Chinese society, one ideal family was patriarchal in authority, that is, young individuals' marriages were decided by their parents and cohabitation was viewed as an abnormal behavior. Due to the great social transformation and sexual liberation in China, individual agency has been on the rise since the 1970s. In this sense, today young individuals prefer to live separately with their parents for more privacy and more freedom of marriage or economic independence. From the rational choice perspective, young people cohabit in order to share living expenditures such as housing and daily meals. Furthermore, with the heavy burden of marriage costs such as a new apartment, a wedding ceremony, and a car, they may choose cohabitation as a prelude to marriage or a trial marriage. Once they decide to get married, they always display more stable marital quality in the future. For those who are capable of living together or who can afford marriage costs, cohabitation can be an option to experience marital life. Regardless of their will to marry, however, the arrangements of one child policy and birth control (Huang 1982; Short et al. 2001) can eventually force individuals who are eager to have a child into legal marriage for the birth permission. Otherwise, those illegally born children are rejected by the household registration system and cannot have access to social welfare.

As for the cohabitation effect, until the widespread utilization of propensity score technique, it is not clear whether the relationship between premarital cohabitation and marital disruption is selective or causal. Lu et al.'s research (2012) merely demonstrated the cohabitation effect in the western context while it is still unclear in oriental cultures. This article's results imply that, similar to Lu el al.'s research conclusion, for those urban individuals married between 1978 and 1999 when cohabitation was not common, premarital cohabitation indeed increased the risk of marital disruption. For those urban individuals married between 2000 and 2010 when cohabitation was prevalent, on the other hand, the cohabitation effect disappeared. Obviously, the causal and diffusive perspectives all apply to the Chinese context. However, the selection effect is not fully decided due to lack of some important covariates and requires further study in the future, even though in this article those observed selective factors partly accounted for subsequent marital dissolution when cohabitation was uncommon.

More specifically, preexisting characteristics of cohabitors affect urban individuals' divorce hazards among urban marriage cohort 1978-1999. However, after those demographic features are balanced by using the propensity score matching method, the cohabitation effect still exists which demonstrates the causal effect. Furthermore, for urban marriage cohort 2000-2010, the Cox models both show that the cohabitation effect no longer appears. One possible reason is that with a high divorce rate in contemporary China, premarital cohabitation is regarded as a test ground for the success of marriage. In this sense, cohabitation plays an important role of excluding unsuitable partners which then could reduce the possibility of subsequent marital disruption.

Since CFPS2010 data do not collect adults' complete information on cohabitation histories, it is highly possible that respondents may have cohabited with multiple partners. In addition, the data also lack partial information on adults' parental social and demographic status such as education, marital status and occupation which could have a negative impact on this article's analysis, but sensitive analysis indicates the conclusion is robust.

Constrained by the data, one possible improvement of this study in the future is to identify the cohabitation history with single cohabitation with the spouse only and serial cohabitation with different partners which could allow scholars to explore the mechanism of how cohabitation affects marital disruption. Another way is to collect as many as possible of those covariates that are associated with cohabitation and marital disruption to improve the robustness of the research conclusion.

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Appendix

Table 1. The Distribution of Cohabitation with the Spouse Prior to First Marriage in Different Marital Periods

Marital periods	*National Sample(N=29,096)*	*Urban Sample(N=8432)*
Before 1978	110 *(1.41)*	43 *(1.82)*
1978-1999	884 (6.01)	326 *(7.56)*
2000-2010	1337 *(20.26)*	500 *(28.36)*
Total	2331 *(8.01%)*	869 *(10.31%)*

Table 2. The Distribution of Duration of Cohabitation in Different Marital Periods

	National Sample (N=2238)		*Urban Sample (N=836)*	
Marital periods	*<=12 months*	*>12months*	*<=12 months*	*>12months*
Before 1978	80 *(76.92)*	24 *(23.08)*	35 *(87.50)*	5 *(12.50)*
1978-1999	732 *(87.35)*	106 *(12.65)*	258 *(83.77)*	50 *(16.23)*
2000-2010	1078 *(83.18)*	218 *(16.82)*	388 *(79.51)*	100 *(20.49)*
Total	1890 *(84.45%)*	348 *(15.55%)*	681 *(81.46%)*	155 *(18.54%)*

Table 3. The Distribution of Divorcers in Different Marital Periods

Marital periods	*National Sample(N=29,270)*	*Urban Sample (N=8478)*
Before 1978	135 *(1.73)*	53 *(2.24)*
1978-1999	526 *(3.57)*	278 *(6.44)*
2000-2010	275 *(4.08)*	112 *(6.23)*
Total	936 *(3.20%)*	443 *(5.23%)*

Table 4. The Descriptive Analysis of Marriage Cohorts 1978-1999 and 2000-2010

Covariates	*Urban Marriage Cohort 1978-1999 (N=4112)*					*Urban Marriage Cohort 2000-2010 (N=1467)*				
	No cohabitation	*Cohabit with the spouse prior to first marriage*	*Absolute Standardized differences*		*Coefficient of logistic regression for predicting propensity score*	*No cohabitation*	*Cohabit with the spouse prior to first marriage*	*Absolute Standardized differences*		*Coefficient of logistic regression for predicting propensity score*
			dx	*dxm*				*dx*	*dxm*	
%Sample (N)	92.19	7.81				66.73	33.27			
% Gender (1=male)	49.54	52.34	0.056	0.064	.27**	49.85	51.64	0.036	0.086	.09
% Education										
Primary school(below)[a]	15.43	12.15	0.095	0.021		4.39	5.33	0.043	0.068	
Junior Mid. School	37.19	44.24	0.144	0.074	.32	25.43	23.98	0.034	0.048	-.28
Senior Mid. School(above)	47.38	43.61	0.076	0.059	.97	70.17	70.70	0.011	0.077	-.30
%Income (*yuan*)										
0-12000[a]	51.33	43.61	0.155	0.036		39.73	33.81	0.123	0.026	
12001-24000	25.90	24.92	0.023	0.016	.06	29.52	24.80	0.106	0.048	.01
24001-60000	19.28	26.17	0.165	0.047	.46**	25.74	31.97	0.138	0.086	.43**
Above 60000	3.48	3.62	0.089	0.021	.49*	5.01	9.43	0.171	0.029	.93***
% Single Child Family	6.65	4.67	0.085	0.158	-.35	27.37	28.48	0.025	0.106	-.19
% CCP Membership (1=yes)	15.64	8.41	0.224	0.020	-.69**	13.79	9.02	0.150	0.055	-.48**
%Living Area of China										
East residence	51.38	57.32	0.119	0.028	.48**	45.97	57.38	0.230	0.071	.61**
Middle residence	36.14	33.33	0.059	0.012	.16	38.71	32.58	0.128	0.084	.29
West residence[a]	12.48	9.35	0.100	0.030		15.32	10.04	0.159	0.014	
%Birth year										
1939-1959(1935-1977)[a]	36.48	14.33	0.526	0.013		38.71	26.84	0.255	0.023	
1960-1969(1978-1984)	42.94	37.69	0.107	0.049	.79***	51.99	60.86	0.179	0.017	.51***
1970-1987(1985-1994)	20.58	47.98	0.602	0.039	1.77***	9.30	12.30	0.096	0.006	.76***
% *Hukou* at 12 (1=urban)	56.61	46.73	0.198	0.002	-.25**	58.12	50.00	0.163	0.070	-.33**
Age of First Marriage	24.75	23.96				26.02	25.65			

Note: *p<0.1, **p<0.05, ***p<0.001; "*a*" represents the reference group; the year of birth in parentheses refers to marriage cohort 2000-2010.

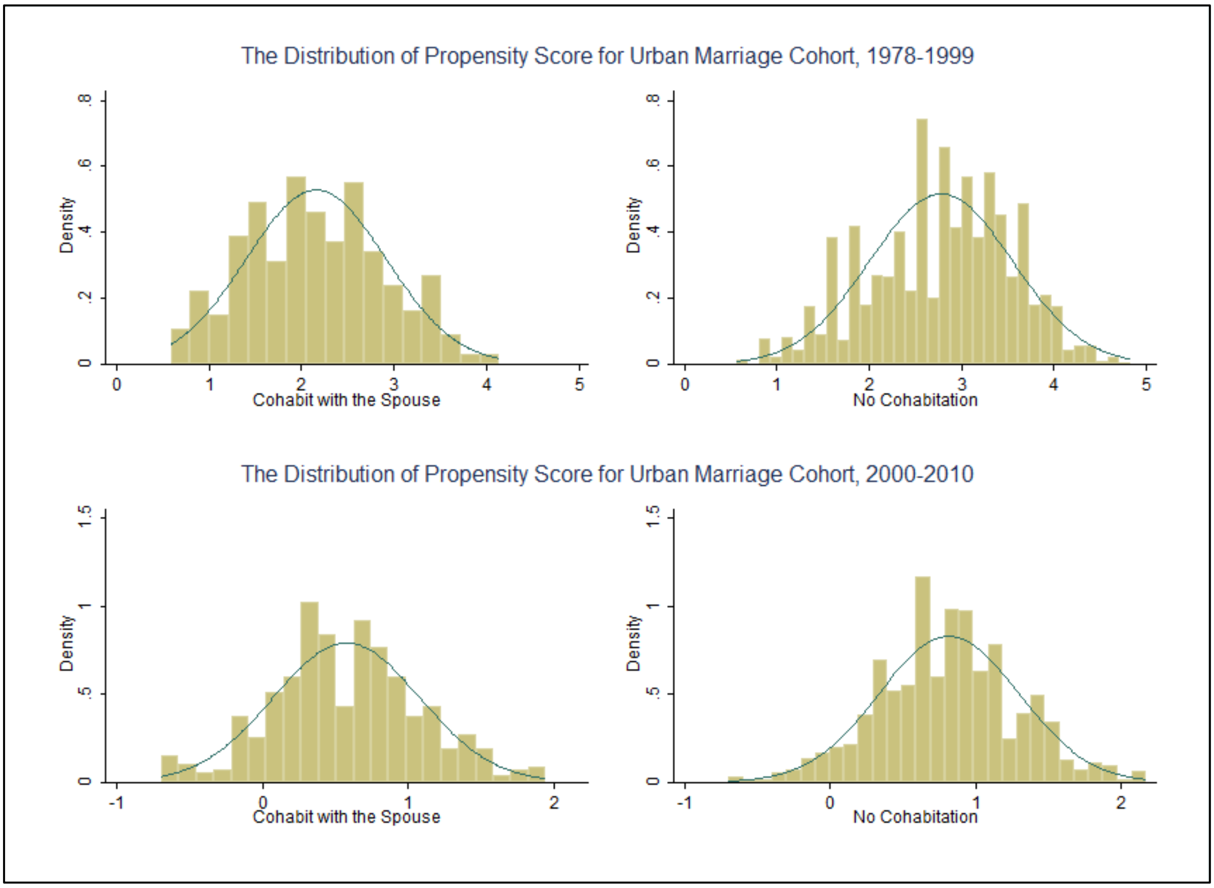

Figure1. The Distribution of Propensity Scores for Two Urban Marriage Cohorts.

Table 5. Divorce Hazard Ratios of Cox Models for Two Marriage Cohorts before Propensity Score Matching

Urban Marriage Cohorts	1978-1999	2000-2010
Cohabitation with the spouse	1.92***	.72
N	*4107*	*1467*
Wald chi2	*88*	*27*

Note: 1) **p*<0.1, ***p*<0.05, ****p*<0.001; 2) to simplify, control variables are not presented here.

Table 6. Divorce Hazard Ratios of Cox Models for Two Marriage Cohorts after Propensity Score Pair Matching

Urban Marriage Cohorts	1978-1999	2000-2010
Cohabitation with the spouse	1.60**	.77
N	*1601*	*976*
Wald chi2	34	59

Note: 1) **p*<0.1, ***p*<0.05, ****p*<0.001; 2) for marriage cohort 1978-1999, optimal pair 1:4 matching, and for marriage cohort 2000-2010, optimal pair 1:1 matching; 3) to simplify, control variables are not presented here.

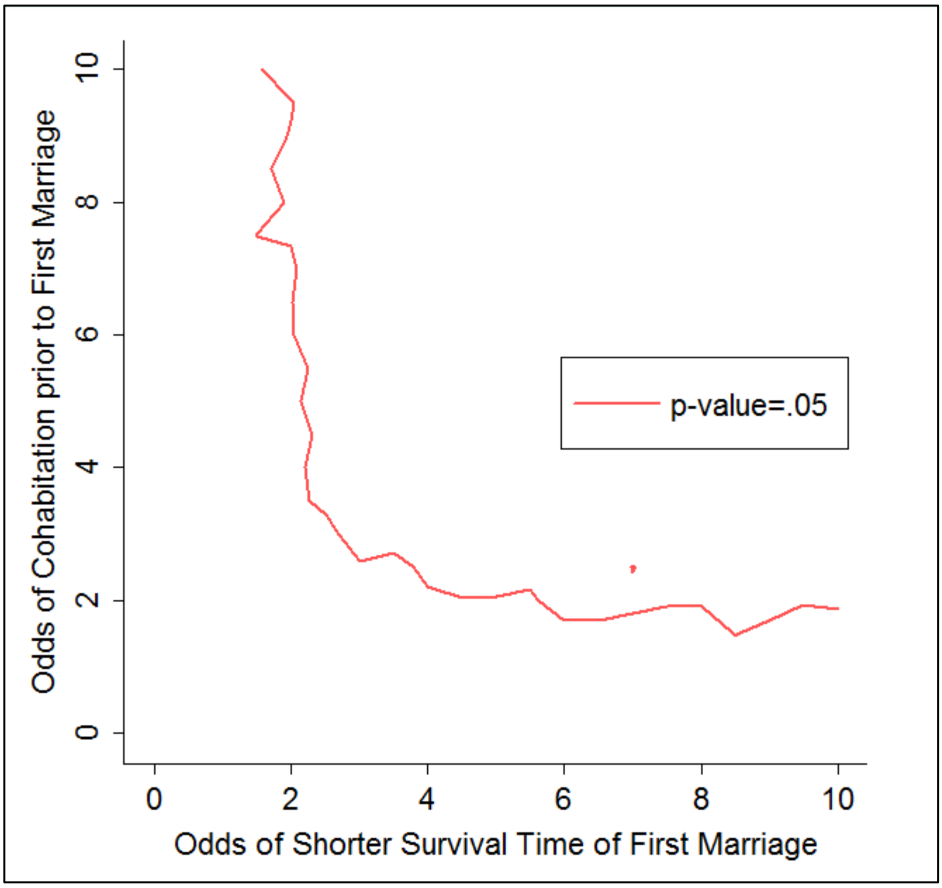

Figure 2. Sensitive Analysis for Marriage Cohort, 1978-1999: Contour Plots of Qualitative Change of Conclusions (*p*=.05)