Health Insurance Coverage and Marriage Behavior: Is There Evidence of Marriage Lock?

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Abstract

Subsidies, taxes, premiums, and eligibility for health insurance may cause "marriage lock," in which couples stay married for the sake of maintaining health insurance coverage. By using the Health and Retirement Study for adults aged 60–70, I examine whether employer-based health insurance coverage for the spouse discourages divorce for spousal health insurance coverage-dependent individuals. Diverse difference-in-difference models provide evidence of a 7% increase in the number of divorces upon achieving Medicare eligibility at age 65 for people with spousal insurance coverage relative to those without it. The estimates thus provide evidence that marriage lock exists.

Keywords: Marriage Lock, Medicare, Employer-sponsored Health Insurance, Marriage Behavior

JEL classification codes: J, I1, D1

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1 Introduction

The predominant source of health insurance in the United States is employer-sponsored health insurance (ESI). Nearly two-thirds of adults under age 65 and three-quarters of all full-time workers obtain health insurance through their employers (U.S. Census Bureau, 2011). A potential cost of this reliance on ESI is the non-portability of insurance across employers, which is likely to result in "job lock," a phenomenon whereby people stay in jobs they might otherwise leave.

A similar concern regarding disruption to health insurance coverage (HIC) may influence the decisions of individuals contemplating divorce. People currently covered by their spouse's ESI lose such coverage on divorce. Potential divorcees may thus face high premiums in the individual health insurance market or the possibly prohibitive health costs of being uninsured. Furthermore, changes in health plans and providers might be disruptive and costly. Unless they have alternative sources of HIC, such as ESI through their own employer or Medicare or Medicaid, this health insurance conundrum could result in "marriage lock," which functions in a similar manner to job lock in that people are forced to remain married despite wanting a divorce.

Individuals dependent on a spouse's ESI prior to divorce are among the most vulnerable to insurance loss after divorce. Such potential divorcees must search for alternative insurance sources to prevent gaps in coverage and may need to rely on the non-group health insurance market. The shortcomings of this market are well documented and can create significant barriers to coverage for divorcees without their own sources of health insurance. From a cost perspective, compared with ESI plans where employers contribute almost 80% of premiums, non-group health insurance policies are more expensive for comparable benefits because the enrollee typically pays the full premium, administrative costs are higher, and coverage is less generous. In 2009, for example, the average annual premium for non-elderly single policies was $2985, while the average premium

for older people was substantially higher (e.g., $5755 for single policies for people aged 60–64).

To obtain an affordable policy, non-group purchasers therefore often forego critical benefits such as pharmaceuticals and mental health services. In addition, in most states, individuals attempting to purchase insurance can be denied coverage because of their health status, age, or other risk factors. If sold a policy, they may be charged more because of these factors, and particular types of care may thus be excluded from their coverage.

A patchwork of federal and state laws has attempted to help dependent spouses obtain HIC after divorce. For example, the Consolidated Omnibus Budget Reconciliation Act, known as COBRA, is a federal law that allows divorcees to continue to use an ex-spouse's coverage for up to 36 months. However, the protection offered under this patchwork has significant limitations. For instance, COBRA's coverage is expensive since enrollees must pay the full cost of the premium (with no premium subsidies) plus a 2% administrative fee, and may thus be out of reach financially for many divorcees.

The underlying theoretical model most commonly used to analyze marriage behavior is based loosely on the Becker model of marriage (Becker, 1981). The Becker model suggests that divorce occurs when the expected utility from being married is less than the expected utility from being single. For people who depend on their spouse's health insurance, leaving a marriage implies leaving the guarantee of subsidized health insurance coverage. The spouse's ESI could thus be considered to be part of the value of the marriage because of the uncertainty of the non-group health insurance marketplace. Hence, many people may decide to stay in their current marriage despite incentives for divorce because they are afraid they may be denied HIC because of pre-existing conditions, lose access to trusted providers, or unable to afford healthcare premiums.

In the United States, most individuals become eligible for public health insurance (Medicare)

at age 65. Attaining Medicare eligibility immediately reduces the value an individual places on a spouse's HIC and therefore on their marriage. According to the Becker model, when an individual who depends on their spouse's HIC qualifies for Medicare at age 65, reductions in the value of marriage increase the probability of divorce. Becker's model thus predicts that individuals whose only source of HIC is through a spouse's insurance plan are more likely to divorce when they first qualify for Medicare than those who have other sources of HIC, suggesting that HIC can indeed serve as a type of marriage lock. Given these concerns, it is surprising that few studies have examined whether the current health insurance system affects marriage behavior.

This study bridges the gap in the body of knowledge on this topic. I examine whether HIC affects late-life divorce by exploiting the abrupt change in HIC that occurs at age 65 (i.e., eligibility for Medicare). By focusing on individuals aged 60–70, the discontinuity in coverage suggests that a difference-in-difference (DID) comparison between the flow of new divorces for individuals dependent on spousal HIC who are younger than 65 and the divorce flow for those who are age 65 and older provides a test of the marriage lock hypothesis.

In particular, I focus on the "divorce flow," or the newly divorced rate (i.e., the rate of new divorces among those currently married). This is distinct from the divorce level (also confusingly called the divorce rate), which represents the proportion of the population currently unmarried due to divorce (a stock concept). Although previous studies have exploited the discontinuity in HIC created by Medicare (e.g., Card, Dobkin, & Maestas, 2008, 2009) to study retirement and insurance decisions, to my knowledge, this is the first study that uses the discontinuity created by Medicare to test the marriage lock hypothesis.

My estimation results support the hypothesis that individuals who lack an alternative source of HIC are more likely to divorce when they become eligible for Medicare than those who have

other sources of HIC. My parameter estimates imply that qualification for Medicare at age 65 increases the probability of divorce by approximately 7% for individuals dependent on spousal insurance coverage compared with those who have other sources of coverage. In addition, I use several triple-difference models to estimate the interaction among spousal employer-based insurance coverage dependence, lack of alternative access to public health insurance (e.g., Medicaid or Medicare obtained before age 65), and age of Medicare eligibility. I find that individuals who have a single access point (i.e., a spousal employer-provided health insurance plan) are approximately 6% more likely to leave their marriage after age 65 than individuals who have access to an alternative source of health insurance prior to age 65. These results are not sensitive to the dependent variables, and I do not find evidence from additional specification estimates that other factors such as retirement or social security are responsible for the increase in divorce flows after an individual turns 65.

The presented results shed light on whether the current health insurance system affects marriage behavior in the United States. They suggest that HIC may serve as a type of marriage lock, possibly because of the high cost of health insurance. When alternative cheap or almost free health insurance plans are available such as Medicare, couples may therefore be more likely to divorce. Further, the results suggest that the emphasis of the United States on ESI may limit the flexibility of marriage and influence the decisions of individuals regarding marriage and divorce; the results are thus of considerable interest to policymakers who promote marriage and marital stability. By contrast, understanding the effects of allowing spousal coverage through ESI and changes in the health insurance markets on marriage behavior is becoming increasing important as the United States continues to restructure its healthcare system.

2 Literature Review

The economics literature on health insurance and family structures has primarily focused on estimating how the marriage and divorce law revolution in the United States has affected marriage behavior and the labor supply of couples as well as the extent to which HIC has influenced labor force participation and self-employment. A large body of work in family economics analyzes how various public policies affect people's marriage behavior and family structure (e.g., the unilateral divorce law and same-sex marriage law). For example, Peters (1986) shows that unilateral divorce has no effect on the probability of divorce as suggested by the Coase theorem, while Allen (1992) argues that transaction cost is significant in marital bargaining, finding that the divorce rate increased significantly once no-fault divorce laws were introduced.

Rasul (2006) and Mechoulan (2006) suggest that the divorce rate rose sharply following the adoption of unilateral laws; however, the increase was reversed within a decade, possibly because of better marital sorting. Gruber (2011) confirms that the implementation of unilateral divorce regulations significantly increased the incidence of divorce by using 40 years of census data to exploit the variation across states and over time in divorce regulation changes. He finds that adults exposed to unilateral divorce regulations as children are less well educated, have lower family incomes, marry earlier, and separate more often. Buchmueller and Carpenter (2010) use the California Health Interview Survey to study the response of same-sex couples to the option of receiving health insurance through a spouse's employer and find that female homosexuals are more likely to have insurance through a spouse's employer and less likely to work full-time.

In contrast to the dearth of research on how insurance affects divorce rates, a large stream of the literature examines the extent to which health insurance influences individuals' labor force participation and self-employment decisions. Historically, health benefits were offered in tight

labor markets as a method of attracting employees (Fronstin, 2006). The rationale was that employees who prefer HIC may be willing to forgo other benefits, job attributes, or wages to obtain employer-provided health coverage (Rosen, 1986). Many economists and health policy experts believe that tying HIC to job status causes people to stay in jobs that they might otherwise leave (i.e., job lock). For example, Madrian (1994) estimates that job lock reduces the voluntary turnover rate of those with ESI by 25%, although this rate was revised downward by subsequent studies (see also Rust & Phelan, 1997). Similarly, Rogowski and Karoly (2000) study the role of health insurance in the retirement decisions of older workers. They use data from the 1992 and 1996 waves of the Health and Retirement Survey to demonstrate that access to post-retirement health insurance has a large effect on retirement. They find that older male workers with retiree health benefit offers are more likely to retire than their counterparts who lose employment-based health insurance upon retirement. Gruber and Madrian (2004) document the distortions to the labor market associated with such a system, including limited job-to-job mobility and distorted retirement decisions. They conclude that health insurance has important effects on both labor force participation and job choice, but whether these effects result in large losses of either welfare or efficiency is unclear.

In addition to the literature on job lock, some recent studies have empirically analyzed the effects of HIC on entrepreneurship and self-employment. Fairlie, Kapur, and Gates (2011) use data from the 1996 to 2006 Current Population Survey to find large, statistically significant results indicating that men and women are less likely to start businesses if they do not have a spouse with employer-based insurance and if there is a family member in bad health. They also focus on the increase in the probability of self-employment when an individual becomes eligible for Medicare and is no longer dependent on employment associated with insurance coverage for access to

guaranteed comprehensive insurance coverage. They find a 13% increase in the probability of owning a business once an individual reaches age 65. Further, the study published by the Urban Institute (2013) estimates that an additional 1.5 million people will launch their own business and become self-employed because of the key provisions in the Affordable Care Act (ACA) that make high-quality insurance on the open market more accessible and affordable. Significant barriers to coverage are eliminated and more people are able to start their own businesses without risking denial of coverage or being unable to afford the premiums.

Similar effects of HIC may also apply to welfare recipients or the disabled population; tying HIC to benefits may exacerbate the strong incentives to never leave welfare/disability. Evidence suggests that such "welfare lock" is statistically significant but relatively small in magnitude (Ellwood & Adams, 1990; Yelowitz, 1995; Livermore, Roche, & Prenovitz, 2009). In addition, access to spousal health insurance has been used in several studies of health insurance and job mobility or business creation (Madrian, 1994; Holtz-Eakin, Penrod, & Rosen, 1996; Kapur, 1998; Madrian & Lefgren, 1998; Wellington, 2001).

3 Model and data

In this section, I present a model for the decision to divorce to understand how HIC affects marriage behavior, especially for potential divorcees approaching age 65 that will qualify for Medicare. Becker, Landes, and Michael (1977) and Becker (1981) suggest that divorce occurs when the expected utility from being married is less than that from being single. This situation exists because marriage as a transaction may be costly to enter and leave in terms of time, money, and effort.

Based on Becker's model on decision-making in marriage, let us first consider a general model with identical men and women that seek each other in the marriage market, with strictly quasi-linear preferences, as follows:

$$U_j = V_j + (H_j - \pi_j), \quad j = M, S. \tag{3.1}$$

where *M* denotes married and *S* denotes single/divorced. V is the utility gain measured in dollar units from a set of variables that could affect the marriage/divorce decision (e.g., children, income, retirement, and love) and H is the utility gain measured in dollar units from having HIC. π is the premium/cost of health insurance and (H − π) is the net value from having HIC.

For simplicity, I assume no variation in insurance quality (i.e., H is assumed to be the same for all insurance plans). The premiums available to divorcees at different ages vary. In addition, I assume individuals only have ESI in marriage and do not change their health insurance choices if they stay married. Divorcees choose health insurance plans on the individual non-group market before age 65 and on the Medicare market thereafter. That is,

$$H_S = H_M,$$

$$\pi_M = \pi_{ESI},$$

$$\pi_S = \begin{cases} \pi_{Nongroup}, if\ age < 65 \\ \pi_{Medicare}, if\ age \geq 65 \end{cases} \text{ and}$$

$\pi_{Medicare} < \pi_{ESI} < \pi_{Nongroup}$.

Figure 1 illustrates health insurance premiums for potential divorcees by age. The figure shows that premiums in the individual non-group market are high and that they keep increasing from age 60 to age 65. Then, after individuals reach age 65, premiums decrease sharply to a very low and constant level because of Medicare.

To decide whether to leave or enter into a marriage, individuals choose between M and S to maximize the following:

$Max\ [U_M - U_S, 0]$

$if\ U_M - U_S \geq 0$, he/she stays married; $if\ U_M - U_S < 0, he/she\ divorce$s.

Next, I have

$$U_M - U_S = V_M + (H_M - \pi_M) - V_S - (H_S - \pi_S) = V_M - V_S + (\pi_S - \pi_M)$$

$$= \begin{cases} V_M - V_S + (\pi_{Nongroup} - \pi_{ESI}), if\ age < 65 \\ V_M - V_S + (\pi_{Medicare} - \pi_{ESI}), if\ age \geq 65 \end{cases} \quad (3.2)$$

In conclusion, before age 65, individuals stay married as long as

$\pi_{Nongroup} \geq \pi_{ESI} + (V_S - V_M)$. After age 65, divorce occurs as long as

$\pi_{Medicare} < \pi_{ESI} + (V_S - V_M)$.

If $V_S - V_M = 0$, that is, the basic utility of being single equals the basic utility of staying married despite HIC, the individual is indifferent between divorce and marriage. Because $\pi_{Medicare} < \pi_{ESI} < \pi_{Nongroup}$, potential divorcees choose to stay married before reaching 65 and divorce thereafter.

Figure 2 shows the decision-making process in Equation (3.2) by illustrating the net utility gain from marriage for potential divorcees as a function of age. Because premiums in the individual non-group market keep increasing from age 60 to age 65, the net utility gains from marriage keep rising, too. However, premiums decrease sharply to the subsidized, constant Medicare premium level after age 65, and the utility gain from marriage also drops sharply. If the net utility gain is still larger than or equal to zero, the model predicts that the individual will choose to stay married; if the net utility gain from marriage is below zero, the divorce incentives increase, possibly causing the individual to choose to divorce.

Finally, I add some randomness to the identical individual model by introducing a random error term ε_{it} to Equation (3.1). Now I have

$$\Delta U_{it} = \Delta V_t + \Delta \pi_t + \varepsilon_{it}, \quad (3.3)$$

where $\Delta U_{it} = U_{Mit} - U_{Sit}$, $\Delta V_t = V_{Mt} - V_{St}$, $\Delta \pi_t = \pi_{St} - \pi_{Mt}$, and $\varepsilon_{it} = \varepsilon_{Mit} - \varepsilon_{Sit}$.

Thus, individuals choose to divorce if $\varepsilon_{it} < -\Delta V_t - \Delta \pi_t$. Note that ΔV_t is assumed to be identical for everyone and $\Delta \pi_t$ depends only on age.

I make two assumptions about the distribution of ε_{it}. On the one hand, if ε_{it} is independent and identically distributed, the probability of getting divorced in each period rises when $\Delta \pi_t$becomes less positive. Thus, this model implies a shift in the divorce curve after individuals become eligible for Medicare. On the other hand, if ε_{it} is a permanent individual effect, there is a spike in the divorce rate at the time of Medicare eligibility.

The real world could be a combination of the transitory and permanent errors. That is, the abrupt change in HIC at age 65 due to Medicare will increase the divorce flow at age 65 as well as shift the divorce pattern after age 65. I therefore expect to see a spike in the divorce flow at age 65 and a shift in the level of divorce flow rates for all ages after age 66.

The basic model can also be applied to marriage decisions and the effect of the health insurance exchange on the divorce and marriage decision under the ACA healthcare reform. In other words, a model allows for changes over time in the utility from marriage and in the available premiums on various health insurance markets.

I use Health and Retirement Study (HRS) data to study whether qualification for Medicare increases late-life divorce flows. The HRS is a longitudinal panel study that surveys a representative sample of more than 26,000 Americans over the age of 50 every two years, collecting information on every respondent's income, work, retirement, marriage status, assets, social security incomes, pension plans, health insurance, disability, health status, and healthcare expenditures. The HRS questions ask individuals whether they divorced between the recent interview waves (i.e., which are every two years); thus, I use a probability distribution for the age

of divorce for each respondent who reported that they divorced between the most recent interview waves. I empirically define the divorce flow in my regressions as a dummy variable equal to one if an individual is newly divorced between these two waves. Furthermore, I use simulations to check for robustness.

4 Methods

Because an effect at the group level might exist (i.e., age clustering), I explore two approaches to control for the potential clustering of errors. I first follow the one-step method to estimate the Eicker–White clustered standard errors at the group level. However, the standard asymptotic arguments for the consistency of clustered standard errors may not apply with the small number of groups in this study's context; hence, I still run the risk of underestimating standard errors and over-rejecting the null hypothesis by using the one-step approach. Therefore, I adopt the two-step estimator suggested by Donald and Lang (2007) and make the generous assumption that unobserved cluster effects are drawn from a homoscedastic normal distribution.

I first use a DID model to examine whether HIC affects divorce rates for individuals with spousal HIC dependence by exploiting the discontinuity created at age 65 when individuals qualify for Medicare. Then, I use the triple-difference approach to study how the variation at the age 65 cutoff has affected individuals who have a single source of spousal HIC versus individuals who have their own public HIC.

I first construct the main experimental group of spousal coverage dependence, in which individuals either provide ESI to their spouse or receive coverage from the spouse's ESI. I then isolate the effects of the "Medicare notch" on late-life divorce by estimating the interaction term between the age eligibility for Medicare and group dummy for individuals with spousal coverage

dependence, addressing concerns about the potential influence of observables such as age, retirement, and social security benefits on the results. The approach is useful for identifying whether marriage lock exists for individuals with spousal coverage dependence. Empirically, I estimate the following model:

$$Y_{ist} = \beta 1 + \beta 2 M_{it} + \beta 3 T_{it} + \beta 4 (M_{it} * T_{it}) + \beta 5 X_{ist} + \lambda_t + \delta_s + \varepsilon_{ist}\,, \qquad (5.1.1)$$

where Y_{ist} equals one if the individual divorced between the interview waves. M_{it} denotes whether an individual or his/her spouse is equal to or older than age 65. T_{it} denotes whether an individual is in the treatment group of spousal coverage dependence, that is, whether the individual provides ESI coverage to or receives it from his/her spouse. The coefficient of the interaction between eligibility for the treatment group and qualification for Medicare at age 65, $\beta 4$, captures the DID estimate for marriage lock. In addition, X is a vector of the demographic and control variables, λ_t is the year dummy, and δ_s is the region effect.

The HRS interviews respondents every 2 years and asks them whether they divorced between recent interview waves. I cannot, however, identify the actual year or age of divorce for individuals who reported that they became divorced between recent interview waves. There are thus three possibilities for the actual year of divorce: people could divorce in the current interview year, the past year, or two years before the interview year (i.e., t, $t-1$, or $t-2$, respectively.) Therefore, I estimate this model by assigning a probability to respondents' divorce years according to the distribution assumption based on the weights of the length of time between interview waves. That is, people could divorce in the past full year ($t-1$) with probability 0.5, in the current interview year (t) with probability 0.25, or two years before the interview year ($t-2$) with probability 0.25. [1]

In addition, to further investigate the Medicare notch effect on late-life divorce for individuals

with spousal coverage dependence as well as other sources of public HIC, I estimate a triple-difference model for individuals who have a single source of spousal HIC compared with those having their own public HIC. Individuals who have only a single source of spousal employer-provided health insurance plan are supposed to be more likely to leave marriage after age 65 than individuals who have access to an alternative source of public health insurance (e.g., Medicaid or Medicare obtained before age 65). Empirically, I estimate the following DID model:

$$Y_{ist} = \beta 1 + \beta 2 M_{it} + \beta 3 T_{it} + \beta 4 G_{it} + \beta 5 (T_{it} * G_{it}) + \beta 6 (M_{is} * G_{it}) + \beta 7 (M_{it} * T_{it} * G_{it}) + \beta 8 X_{ist} + \lambda_t + \delta_s + \varepsilon_{ist} \quad (5.1.2)$$

where Y_{ist}, M_{it}, and T_{it} are as before and G_{it} denotes whether an individual has other public health insurance such as Medicaid or Medicare obtained before age 65. The coefficient of the interaction term among eligibility for the spousal coverage dependence group, the group dummy for owning other public health insurance, and qualification for Medicare at age 65, $\beta 7$, captures the DID estimate of marriage lock. In addition, as before, X is a vector of the demographic and control variables, λ_t is the year dummy, and δ_s is the region effect.

5 Results

In this study, I take advantage of the abrupt change in HIC occurring at age 65 because of Medicare eligibility to explore whether this gain in health insurance encourages individuals with spousal coverage dependence to divorce. To focus the analysis around this age cutoff, I restrict the sample to individuals aged 60–70 who are either married or divorced. As the number of observations falls and coefficient estimates become erratic when the age of the older spouse exceeds 71, this age is set to be between 58 and 71. By using the DID and triple-difference estimations, I find that individuals who depend on spousal HIC are more likely to divorce upon

achieving Medicare eligibility at age 65 than those without it.

DID Estimation

I cannot obtain the direct effect of HIC on divorce from the whole population because such an effect may be contaminated by unmeasured variables (e.g., marriage and job quality). Therefore, I only focus on individuals with spousal coverage dependence whose divorce decisions may be affected by HIC. Because individuals with HIC dependence no longer have to be concerned about losing spousal HIC after age 65, the value they place on spousal HIC or current marriage is reduced. According to the model, the probability of divorce should increase after age 65 for these individuals.

Table 1 reports the DID estimates from Equation (5.1.1) considering whether either spouse is age 65 or older. I report both the one-step OLS estimates and the two-step estimates. The coefficient of the interaction term between the age 65 cutoff dummy variable and spousal coverage dependence group dummy is positive and statistically significant under both the one-step OLS and the two-step estimations, suggesting that individuals with spousal coverage dependence are approximately 7% more likely than individuals without such dependence to divorce when either of the spouses qualifies for Medicare at age 65. In other words, individuals with spousal coverage dependence are more likely to be deterred from divorce before age 65 because of their current health insurance status. The positive and significant coefficient is consistent with the notion that a spouse's employer-provided HIC is a disincentive to divorce before age 65. Generally, the signs, magnitudes, and significance levels of the coefficients are stable across specifications. Further, the divorce rates decrease with the number of children, years married, times married, age, and family income, whereas personal income, education level, disability, and retirement increase the divorce

rate.

I also investigate whether the effect of Medicare eligibility on late-life divorce is a one-time effect at age 65 or a permanent effect that persists after age 65. To do so, I create two age cutoff dummy variables for Medicare eligibility: one where either spouse's age is equal to 65 (age = 65) and the other where either spouse is older than 66 (age $\geq$ 66). Table 2 reports the DID estimates from Equation (5.1.1) using these two age cutoff dummies. The coefficients of the age = 65 and age $\geq$ 66 dummies suggest that individuals are approximately 8% and 6% more likely to divorce at age 65 when they qualify for Medicare, respectively. The coefficients of the interaction term between the age = 65 dummy and spousal coverage dependence group dummy are significant in both the one-step OLS estimation and the two-step estimation, while the interaction term evolving the age $\geq$ 66 dummy is not significant in the two-step estimation.

Figure 3 depicts the age variation in the divorce flow between the ESI coverage dependence group and "no such dependence" group by plotting the difference in the first-step coefficients between the groups. [2] Figure 3 shows a spike at age 65, which suggests that many individuals with spousal coverage dependence divorce at age 65. In addition to the high premiums and cost sharing on the non-group market, potential divorcees choose to stay married because they are afraid of being rejected from new insurance policies after their divorce owing to pre-existing health conditions. However, COBRA allows divorcees to stay on their ex-spouse's ESI coverage for up to three years by paying 102% of the full premium themselves, which is nevertheless more affordable than the plans on the non-group market. Under this arrangement, divorcees will furthermore not be rejected for coverage based on pre-existing conditions. The existence of COBRA's policies reduces the cost of divorce as people approach age 65, and so Figure 3 also shows a build-up starting from age 62 rather than a perfect spike at age 65.

In addition, consistent with the results in Figure 3, the econometrics presented in Table 3 show that 65 is the most important age. Table 3 reports the results of a test of the spike and shift in Figure 3 as well as a placebo test for other ages, which regress the difference in the first-step coefficients between two groups on the age trend, an age dummy for age X (X=58, 59…71), and an age ≥ 66 dummy. Specification 8 including a dummy for age 65 is the key regression, which tests for the spike at age 65 and the shift after age 65. All other specifications are placebo tests for other ages. The results in Table 3 show that from age 58 to 71, only the coefficient for the age 65 dummy is significant and has the largest effect (about 8%). The coefficients of the dummies for other ages are not significant and are much smaller in magnitude.

The findings drawn from Figure 3 and Table 3 confirm the theoretical prediction of a spike in the divorce flow at age 65 for individuals with spousal coverage dependence. Figure 3 also shows a higher divorce flow after age 65 than before, although the estimated coefficient for the age ≥ 66 dummy is not statistically significant (see Table 3).

Triple-difference Estimation

Some individuals with spousal coverage dependence may also have other public health insurance for themselves such as Medicaid or Medicare obtained before age 65 because of disabilities. Individuals with their own sources of public health insurance are supposed to be less dependent on spousal coverage and thus less affected by marriage lock. While individuals with spousal coverage dependence face a potential disruption in HIC when leaving their current marriage, individuals with their own public HIC may not. Thus, individuals who rely on their spouse's HIC and do not have access to an alternative plan may be more likely to be deterred from divorce because of HIC issues before age 65.

Therefore, I use the triple-difference model in Equation (5.1.2) to estimate the interaction among the age 65 cutoff dummy variable, spousal ESI coverage dependence group dummy, and "lacking other public HIC" group dummy. The "lacking other public HIC" group is defined as individuals who do not have Medicaid or "pre-65" Medicare. The coefficients of the interaction term shown in Table 4 are positive and statistically significant, suggesting that among individuals with spousal ESI coverage dependence, those with a single source of spousal ESI coverage are approximately 6% more likely to divorce when they qualify for Medicare at age 65 than those with other public health insurance such as Medicaid or Medicare obtained before age 65. These positive and significant estimated effects are robust for all specifications, which suggests that a lack of access to one's own health insurance is a disincentive to divorce before age 65 for those with spousal coverage dependence.

Table 5 reports the triple-difference estimates from Equation (5.1.2) using both "whether either spouse's age is equal to age 65" and "whether either spouse's age is older than age 66" as the age cutoff dummies for Medicare qualification. The coefficients of the interaction terms of the age 65 cutoff dummy variable, spousal ESI coverage dependence group dummy, and "lacking other public HIC" group dummy are also positive and statistically significant, suggesting that among individuals with spousal ESI coverage dependence, those that only have spousal ESI coverage are approximately 8% more likely to divorce when they qualify for Medicare at age 65 than people with other public health insurance such as Medicaid or Medicare obtained before age 65. Furthermore, the estimates show that they are approximately 5% more likely to divorce after age 65.

Potentially Confounding Factors

The changes in the probability of divorce observed around age 65 may be due to other changes in work status or social security benefits, which may be an analytical concern. For instance, individuals may divorce at age 65 because of their transition into retirement, which may be irrelevant to qualifying for Medicare. Thus, I investigate whether other confounding factors cause changes in marriage behavior around age 65 by including controls for retirement and social security into my regressions.

First, the estimates of the coefficients of retirement are reported in all specifications (Tables 1–5). I find positive coefficient estimates for the retirement variable, which are not significant for most specifications; however, the key coefficient estimates of the interaction term between the age cutoff for Medicare and spousal coverage dependence group remain significant and robust, [3] which suggests that retirement is not responsible for the primary changes in marriage behavior at age 65. Indeed, the effect of Medicare coverage on late-life divorce may be underestimated because individuals who retire because they qualify for Medicare at age 65 may decide to divorce later because of problems and conflicts occurring after retirement. Thus, this kind of divorce flow may be attributed to the effect of Medicare eligibility at age 65 on late-life divorce.

Second, the coefficients of social security income reported in all specifications of Tables 1–5 are negative and insignificant, suggesting that individuals may be less likely to divorce when they have higher social security benefits. In summary, the addition of the covariates does not have a significant effect on the estimated relationship between the key interaction term and divorce flow. The coefficient estimates of the interaction term between the age cutoff for the Medicare dummy and spousal coverage dependence group dummy remain significant and robust, which rules out the possibility that retirement or social security benefits generate the main change in marriage behavior around age 65.

Simulation for Robustness Check

One limitation of using panel data from the HRS is the reliance on the distribution assumptions for respondents' ages of divorce. Thus, I use a simulation to run the probability assignment process 10,000 times. I find roughly similar-sized point estimates, as shown in Table 6, which reports the simulated results for both the one-step OLS and the two-step estimations for both the DID and triple-difference models, which are consistent with the previous estimation results. Panel A reports the coefficient and standard errors for the key independent variable, which is the interaction term between the age = 65 cutoff dummy variable and spousal coverage dependence group dummy in the DID estimations or the interaction term among the age = 65 cutoff dummy variable, spousal coverage dependence group dummy, and "without other public health insurance" group dummy in the triple-difference estimations. Similarly, Panel B reports the coefficient and standard errors for the key independent variable, which is the interaction term between the two age cutoff dummies (i.e., age = 65 and age $\geq$ 66) and the spousal coverage dependence group dummy in the DID estimations and the interaction term among the two cutoff dummy variables, spousal coverage dependence group dummy, and "without other public health insurance" group dummy in the triple-difference estimations. In general, the simulated estimates in Table 6 show robustness and the results do not appear to be sensitive to changes in covariates or estimation methods.

All the estimations rest on the assumption that potential divorcees rely heavily on Medicare coverage rather than on other possible sources of HIC after divorce. The best protection against insurance loss for those individuals is stable long-term employment in jobs that offer a direct source of insurance coverage. Although some spouses may actively look for jobs with health insurance during a divorce, it is unlikely that this search drives the entire relationship, especially

in late adulthood. I do not, however, capture the effect from divorcees who concurrently find insured jobs during or after the divorce.

6 Discussion and conclusion

This study examines whether employer-based HIC for the spouse discourages divorce for spousal HIC-dependent individuals. The parameter estimates presented herein imply that age eligibility for Medicare among married couples aged 60 to 70 with spousal coverage dependence increases the probability of divorce by 7%. I also find that divorce flow rates at age 65, when people qualify for Medicare, are substantially lower among those who have their own public insurance compared with those who have insurance coverage only through a spouse. My estimates thus provide some evidence of marriage lock. These estimates further suggest that HIC could serve as a type of marriage lock and that the price of health insurance could be a key factor in this regard.

The enacted ACA stipulates that individuals are able to purchase insurance from insurance exchanges. Insurers are not allowed to have pre-existing condition exclusions or premiums priced on the basis of health status. These features of the ACA may encourage marriage flexibility by providing potential divorcees with a health insurance option should they leave their current marriage. However, the disparity between the value of health coverage through insurance exchanges and the value of coverage through existing employer plans is likely to persist for some time (Eibner et al., 2010).

With the deepening of the healthcare reform, health insurance will become increasingly affordable. In July 2013, New York State insurance regulators said that individuals buying health insurance on their own would see their premiums drop in the next year as the changes under the federal healthcare law take effect. In addition, they approved rates for 2014 that are at least 50%

lower on average than those currently available in New York. The extraordinary decline in New York's insurance rates for individual consumers demonstrates the future direction of health insurance exchanges such as universal coverage and more affordable individual health insurance plans, which will lead to a reduction in marriage lock. Hence, an interesting area for future research would be to investigate the impact of these changes on health insurance markets and marriage behavior under the healthcare reform. For instance, it will be important to examine whether marriage lock disappears when insurance prices in exchange markets are sufficiently low and whether an equilibrium exists, for example, when the price in the exchange market equals the price people pay for ESI coverage. Indeed, as more states such as New York set up exchanges under the ACA, where competition and transparency in exchange marketplaces are leading to more affordable HIC, marriage lock is likely to eventually disappear as the price of individual health insurance decreases.

Table 1: DID Estimates of the Divorce Rate for Medicare Eligibility (Age ≥ 65)

Whether the Individual Is Recently Divorced	OLS		Two-step Estimator	
	(1)	(2)	(3)	(4)
Either Spouse's Medicare Eligibility (Age ≥ 65)	-.0539***	-.0581***	-.0616***	-.0469***
	(.0123)	(.0123)	(.0106)	(.0127)
Spousal Coverage Dependence Group	-.0701***	-.0123***	-.0701***	-.3265**
	(.0032)	(.0019)	(.0067)	(.1364)
Spousal Coverage Dependence *Either Spouse's (Age ≥ 65)	.0716***	.0708***	.0695***	.0400**
	(.0082)	(.0078)	(.0094)	(.0180)
Personal Income	3.84e-	3.90e-	3.86e-07***	3.86e-07***
	(1.07e-07)	(1.04e-07)	(8.10e-08)	(8.10e-08)
Family Income	-1.76e-	-1.70e-	-1.76e-	-1.76e-
	(6.62e-08)	(6.28e-08)	(2.61e-08)	(2.61e-08)
Education Level	.0017*	.0019*	.0017*	.0017*
	(.0009)	(.0009)	(.0010)	(.0010)
Self-Reported Health Status	.0036	.0033	.0037	.0037
	(.0025)	(.0025)	(.0026)	(.0026)
Gender	.0453***	.0425***	.0448***	.0448***
	(.0059)	(.0058)	(.0059)	(.0059)
Race	-.0036	-.0029	-.0037	-.0037
	(.0076)	(.0075)	(.0053)	(.0053)
Disability	.0156	.0151	.0149	.0149
	(.0089)	(.0088)	(.0092)	(.0092)
Years Married	-.0065***	-.0062***	-.0064***	-.0064***
	(.0002)	(.0002)	(.0002)	(.0002)
Times Married	-.0120**	-.0096*	-.0118***	-.0118***
	(.0053)	(.0051)	(.0038)	(.0038)
Number of Children	-.0081***	-.0080***	-.0080***	-.0080***
	(.0006)	(.0006)	(.0013)	(.0013)
Post-Retirement ESI	-.0023	-.0022	-.0020	-.0020
	(.0065)	(.0066)	(.0062)	(.0062)
Age	-.0075***	-.0073***	-.0096***	-.0117***
	(.0016)	(.0016)	(.0012)	(.0015)
Retirement	.0068	.0070	.0070	.0070
	(.0050)	(.0050)	(.0045)	(.0045)
Social Security Income	-.0033	-.0043	-.0035	-.0035
	(.0034)	(.0032)	(.0080)	(.0080)
Year Effect & Cohort Effect	Yes	Yes	Yes	Yes
Group-specific Age Trend	No	Yes	No	Yes

Note: Source: HRS 1992–2010. N=57,480. Age range is 60 to 70. Individuals in the sample are either married or divorced. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. Estimates are made under the assumption of probabilities assigned for the age of divorce. Standard errors for OLS regressions are robust,

clustered by age, and shown in parentheses. For the specification of the two-step estimator, the estimates for the first three key independent variables are reported from the second step and all other estimates are reported from the first step.

Table 2: DID Estimates of the Divorce Rate for Medicare Eligibility (Age = 65) and (Age ≥ 66)

Whether the Individual Is Recently Divorced	OLS		Two-step Estimator	
	(1)	(2)	(3)	(4)
Either Spouse's Medicare Eligibility (Age = 65)	-.0597***	.0613***	-.0674***	-.0551***
	(.0137)	(.0143)	(.0148)	(.0140)
Either Spouse's Medicare Eligibility (Age ≥ 66)	-.0501***	-.0559***	-.0612***	-.0381**
	(.0129)	(.0128)	(.0124)	(.0142)
Spousal Coverage Dependence	-.0701***	-.0113***	-.0701***	-.4451***
	(.0033)	(.0019)	(.0069)	(.1464)
Spousal Coverage Dependence *Either Spouse's (Age = 65)	.0815***	.0841***	.0821***	.0575***
	(.0128)	(.0135)	(.0194)	(.0198)
Spousal Coverage Dependence *Either Spouse's (Age ≥ 66)	.0672***	.0749***	.0673***	.0212
	(.0082)	(.0087)	(.0101)	(.0201)
Personal Income	3.84e-	3.90e-	3.86e-07***	3.86e-07***
	(1.07e-07)	(1.04e-07)	(8.10e-08)	(8.10e-08)
Family Income	-1.76e-	-1.70e-	-1.76e-	-1.76e-
	(6.61e-08)	(6.27e-08)	(2.61e-08)	(2.61e-08)
Education Level	.0017*	.0019*	.0017*	.0017*
	(.0009)	(.0009)	(.0010)	(.0010)
Self-Reported Health Status	.0036	.0033	.0037	.0037
	(.0025)	(.0025)	(.0026)	(.0026)
Gender	.0454***	.0425***	.0448***	.0448***
	(.0058)	(.0057)	(.0059)	(.0059)
Race	-.0036	-.0030	-.0037	-.0037
	(.0076)	(.0075)	(.0053)	(.0053)
Disability	.0159	.0151	.0149	.0149
	(.0090)	(.0088)	(.0092)	(.0092)
Years Married	-.0065***	-.0062***	-.0064***	-.0064***
	(.0002)	(.0002)	(.0002)	(.0002)
Times Married	-.0120**	-.0096*	-.0118***	-.0118***
	(.0053)	(.0051)	(.0038)	(.0038)
Number of Children	-.0081***	-.0080***	-.0080***	-.0080***
	(.0005)	(.0005)	(.0013)	(.0013)
Post-Retirement ESI	-.0023	-.0022	-.0020	-.0020
	(.0065)	(.0066)	(.0062)	(.0062)
Age	-.0076***	-.0074***	-.0095***	-.0126***
	(.0016)	(.0016)	(.0013)	(.0017)
Retirement	.0069	.0070	.0070	.0070
	(.0051)	(.0050)	(.0045)	(.0045)
Social Security Income	-.0033	-.0042	-.0035	-.0035
	(.0034)	(.0032)	(.0080)	(.0080)
Year Effect & Cohort Effect	Yes	Yes	Yes	Yes

Group-specific Age Trend	No	Yes	No	Yes

Note: Source: HRS 1992–2010. N=57,480. Age range is 60 to 70. Individuals in the sample are either married or divorced. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. Estimates are made under the assumption of probabilities assigned for the age of divorce. Standard errors for OLS regressions are robust, clustered by age, and shown in parentheses. For the specification of the two-step estimator, the estimates for the first three key independent variables are reported from the second step and all other estimates are reported from the first step.

Table 3: Estimates of the Difference of the First-step Coefficients for All Ages

Difference in the First-step Coefficients	(1) X= 58	(2) X= 59	(3) X= 60	(4) X= 61	(5) X= 62	(6) X= 63	(7) X= 64
Age X Dummy	.0197	-.01795	-.0262	-.0412	-.0083	-.0020	.0182
	.0368	.0343	.0323	.0300	.0326	.0334	.0343
Age ≥ 66 Dummy	-.0130	-.0043	-.0060	-.0101	-.0089	-.0079	.0003
	.0335	.0323	.0313	.0296	.0327	.0338	.0349
Age Trend	.0086*	.0066	.0067	.0071*	.0075*	.0074*	.0066
	.0045	.0042	.0039	.0036	.0040	.0040	.0042
Adjusted R Squared	.3912	.3904	.4123	.0473	.0777	.3739	.3910
Number of Observations	14	14	14	14	14	14	14

Difference in the First-step Coefficients	(8) X= 65	(9) X= 66	(10) X= 67	(11) X= 68	(12) X= 69	(13) X= 70	(14) X= 71
Age X Dummy	.0799***	.0443	.0345	-.0112	-.0037	-.0319	-.0424
	.0254	.0312	.0311	.0333	.0335	.0327	.0332
Age ≥ 66 Dummy	.0353	-.0323	-.0200	-.0048	-.0069	-.0077	-.0128
	.0280	.0318	.0323	.0329	.0324	.0308	.0302
Age Trend	.0027	.0097**	.0083*	.0073*	.0074*	.0082*	.0092**
	.0033	.0037	.0038	.0039	.0040	.0039	.0039
Adjusted R Squared	.6582	.5169	.4459	.3806	.3745	.4282	.4617
Number of Observations	14	14	14	14	14	14	14

Note: Source: HRS 1992–2010. Individuals in the sample are either married or divorced. The spousal coverage dependence group refers to individuals who or whose spouses have a single source of health insurance from the spouse's ESI coverage. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. "The first-step coefficient for the dependent group" is the estimated coefficient of the interaction terms between the treatment group (ESI coverage dependence group) dummy and age (the older age of the spouses) by using the two-step estimation method. "The first-step coefficient for the nondependent group" is defined similarly for the control group without such ESI coverage dependence. The "difference in the first-step coefficients for the dependent and nondependent groups" refers to the difference in the first-step coefficient estimates for the two groups, which is the dependent variable in the regressions. Independent variables include age trend, an age dummy for age X (X=58, 59,…,71), and an age ≥ 66 dummy. Specification 8, including a dummy for age 65, is the key regression and all the other specifications are placebo tests for other ages.

Table 4: Triple-difference Estimates of the Divorce Rate for Medicare Eligibility (Age ≥ 65)

Whether the Individual Is Recently	OLS		Two-step Estimator	
Divorced	(1)	(2)	(3)	(4)
Either Spouse's Medicare	-.0027	-.0028	-.0029	-.0029
Eligibility (Age ≥ 65)	(.0156)	(.0149)	(.0145)	(.0143)
Spousal Coverage Dependence Group	-.0062*	-.0062**	-.0065	-.0076
	(.0034)	(.0021)	(.0091)	(.0129)
Having Neither Medicaid nor	.0684***	.0746***	.0791***	.0741***
Age 65	(.0162)	(.0154)	(.0120)	(.0124)
Spousal Coverage	.0761***	.0631***	.0766***	.0567**
(Age ≥ 65)* Having No Medicaid or	(.0104)	(.0082)	(.0182)	(.0228)
Personal Income	3.73e-	3.79e-	3.75e-	3.75e-
	(1.04e-07)	(1.01e-07)	(8.10e-08)	(8.10e-08)
Family Income	-1.75e-	-1.68e-	-1.75e-07**	-1.75e-07**
	(6.62e-08)	(6.26e-08)	(2.61e-08)	(2.61e-08)
Education Level	.0015	.0017*	.0015	.0015
	(.0009)	(.0009)	(.0010)	(.0010)
Self-Reported Health Status	.0042	.0040	.0044*	.0044*
	(.0025)	(.0025)	(.0026)	(.0026)
Gender	.0456***	.0425***	.0449***	.0449***
	(.0061)	(.0060)	(.0059)	(.0059)
Race	-.0029	-.0022	-.0032	-.0032
	(.0075)	(.0074)	(.0053)	(.0053)
Disability	.0281**	.0284**	.0269*	.0269*
	(.0093)	(.0093)	(.0097)	(.0097)
Years Married	-.0064***	-.0061***	-.0064***	-.0064***
	(.0002)	(.0002)	(.0002)	(.0002)
Times Married	-.0115*	-.0090	-.0113***	-.0113***
	(.0005)	(.0051)	(.0038)	(.0038)
Number of Children	-.0079***	-.0079***	-.0078***	-.0078***
	(.0005)	(.0005)	(.0013)	(.0013)
Post-Retirement ESI	-.0015	-.0014	-.0012	-.0012
	(.0064)	(.0065)	(.0062)	(.0062)
Age	-.0074***	-.0073***	-.0071***	-.0085***
	(.0016)	(.0016)	(.0016)	(.0019)
Retirement	.0073	.0076	.0076*	.0076*
	(.0049)	(.0048)	(.0045)	(.0045)
Social Security Income	-.0019	-.0028	-.0019	-.0019
	(.0037)	(.0035)	(.0080)	(.0080)
Fully Retirement	-.0023	-.0024	-.0022	-.0022
	(.0023)	(.0025)	.(0039)	.(0039)

Year Effect & Cohort Effect	Yes	Yes	Yes	Yes
Group-specific Age Trend	No	Yes	No	Yes

Note: Source: HRS 1992–2010. N=57,480. Age range is 60 to 70. Individuals in the sample are either married or divorced. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. Estimates are made under the assumption of probabilities assigned for the age of divorce. Standard errors for OLS regressions are robust, clustered by age, and shown in parentheses. All specifications also include controls for interaction terms among the Having Neither Medicaid nor Medicare before Age 65 group dummy, Spousal Coverage Dependence Group dummy, and Either Spouse's Medicare Eligibility age dummy. For the specification of the two-step estimator, the estimates of the first four key independent variables are reported from the second step and all other estimates are reported from the first step.

Table 5: Triple-difference Estimates of the Divorce Rate for Medicare Eligibility (Age = 65) and (Age ≥ 66)

Whether the Individual Is Recently Divorced	OLS		Two-step Estimator	
	(1)	(2)	(3)	(4)
Either Spouse's Medicare Eligibility (Age = 65)	-.0111	-.0122	-.0027	-.0083
	(.0205)	(.0193)	(.0147)	(.0145)
Either Spouse's Medicare Eligibility (Age ≥ 66)	-.0095	-.0097	-.0069	-.0071
	(.0152)	(.0147)	(.0169)	(.0167)
Spousal Coverage Dependence	-.0221*	-.0267*	-.0212	-.0194
	(.0121)	(.0136)	(.0137)	(.0135)
Having Neither Medicaid nor Age 65	.0683***	.0745***	.0800***	.0746***
	(.0161)	(.0153)	(.0123)	(.0127)
Spousal Coverage Dependence* (Age = 65)* Having No Medicaid	.0871***	.0796***	.0889**	.0764**
	(.0158)	(.0145)	(.0375)	(.0379)
Spousal Coverage Dependence* (Age ≥ 66)* Having No Medicaid nor	.0710***	.0599***	.0745***	.0512**
	(.0098)	(.0076)	(.0195)	(.0248)
Personal Income	3.73e-	3.79e-	3.75e-	3.75e-
	(1.04e-07)	(1.01e-07)	(8.10e-08)	(8.10e-08)
Family Income	-1.75e-	-1.68e-	-1.75e-07**	-1.75e-07**
	(6.62e-08)	(6.26e-08)	(2.61e-08)	(2.61e-08)
Education Level	.0015	.0017	.0015	.0015
	(.0019)	(.0009)	(.0010)	(.0010)
Self-Reported Health Status	.0043	.0040	.0044*	.0044*
	(.0026)	(.0025)	(.0026)	(.0026)
Gender	.0456***	.0426***	.0449***	.0449***
	(.0060)	(.0059)	(.0059)	(.0059)
Race	-.0029	-.0022	-.0032	-.0032
	(.0075)	(.0074)	(.0053)	(.0053)
Disability	.0281**	.0283**	.0269*	.0269*
	(.0094)	(.0093)	(.0097)	(.0097)
Years Married	-.0064***	-.0061***	-.0064***	-.0064***
	(.0002)	(.0002)	(.0002)	(.0002)
Times Married	-.0115**	-.0090	-.0113***	-.0113***
	(.0053)	(.0051)	(.0038)	(.0038)
Number of Children	-.0080***	-.0079***	-.0078***	-.0078***
	(.0005)	(.0005)	(.0013)	(.0013)
Post-Retirement ESI	-.0016	-.0014	-.0012	-.0012
	(.0064)	(.0064)	(.0062)	(.0062)
Age	-.0075***	-.0073***	-.0076***	-.0092***
	(.0016)	(.0016)	(.0019)	(.0021)

Retirement	.0073	.0075	.0076*	.0076*
	(.0049)	(.0048)	(.0045)	(.0045)
Social Security Income	-.0018	-.0027	-.0019	-.0019
	(.0037)	(.0035)	(.0080)	(.0080)
Fully Retirement	-.0023	-.0025	-.0026	-.0026
	(.0024)	(.0026)	.(0049)	.(0049)
Year Effect & Cohort Effect	Yes	Yes	Yes	Yes
Group-specific Age Trend	No	Yes	No	Yes

Note: Source: HRS 1992–2010. N=57,480. Age range is 60 to 70. Individuals in the sample are either married or divorced. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. Estimates are made under the assumption of probabilities assigned for the age of divorce. Standard errors for OLS regressions are robust, clustered by age, and shown in parentheses. All specifications also include controls for interaction terms among the Having Neither Medicaid nor Medicare before Age 65 group dummy, Spousal Coverage Dependence Group dummy, and Either Spouse's Medicare Eligibility age dummy. For the specification of the two-step estimator, the estimates of the first six key independent variables are reported from the second step and all other estimates are reported from the first step.

Table 6: Simulation Results of the Regression Discontinuity Estimates for Medicare Eligibility

Panel A: Either Spouse'	Diff-in-		Diff-in-Diff (2-		Triple-Diff		Triple-Diff (2-	
Eligibility (Age ≥ 65)	Mean	Std.	Mean	Std.	Mean	Std.	Mean	Std.
Coef. for Key	.0672	.0014	.0651	.0013	.0588	.0012	.0514	.0012
Std. Err. for Key	.0071	.0004	.0280	.0081	.0077	.0004	.0215	.0014
Group-specific Age	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Panel B: Either	Diff-in-		Diff-in-Diff (2-		Triple-Diff		Triple-Diff (2-	
Eligibility (Age = 65)	Mean	Std.	Mean	Std.	Mean	Std.	Mean	Std.
Coef. for Key	.0718	.0035	.0735	.0031	.0778	.0036	.0732	.0032
Std. Err. for Key	.0095	.0012	.0245	.0076	.0092	.0012	.0314	.0047
Coef. for Key	.0612	.0012	.0603	.0018	.0567	.0012	.0496	.0048
Std. Err. for Key	.0066	.0004	.0512	.0094	.0064	.0004	.0229	.0038
Group-specific Age	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Source: HRS 1992–2010. N=57,480. Simulation times: 10,000. Age range is 60 to 70. Individuals in the sample are either married or divorced. Standard errors for OLS regressions are robust and clustered by age. The regressions are based on the assumption that people could divorce in the full past year with probability 0.5, in the current interview year with probability 0.25, or in two years before the interview year with probability 0.25. The dependent variable is a dummy variable that equals one if the individual divorced between the interview waves for all regressions; the main independent variables are the interaction terms among the Having Neither Medicaid nor Medicare before Age 65 group dummy, Spousal Coverage Dependence Group dummy, and Either Spouse' Medicare Eligibility age dummies, with coefficients and standard errors reported in the table. Other important control variables include age, income, education, gender, race, disability, years married, times married, number of children, health status, retirement, private HIC, and social security benefits. All regressions control for the year and region fixed effects as well as spousal coverage group-specific age trend.

Figure 1: Health Insurance Premiums for Potential Divorcees by Age

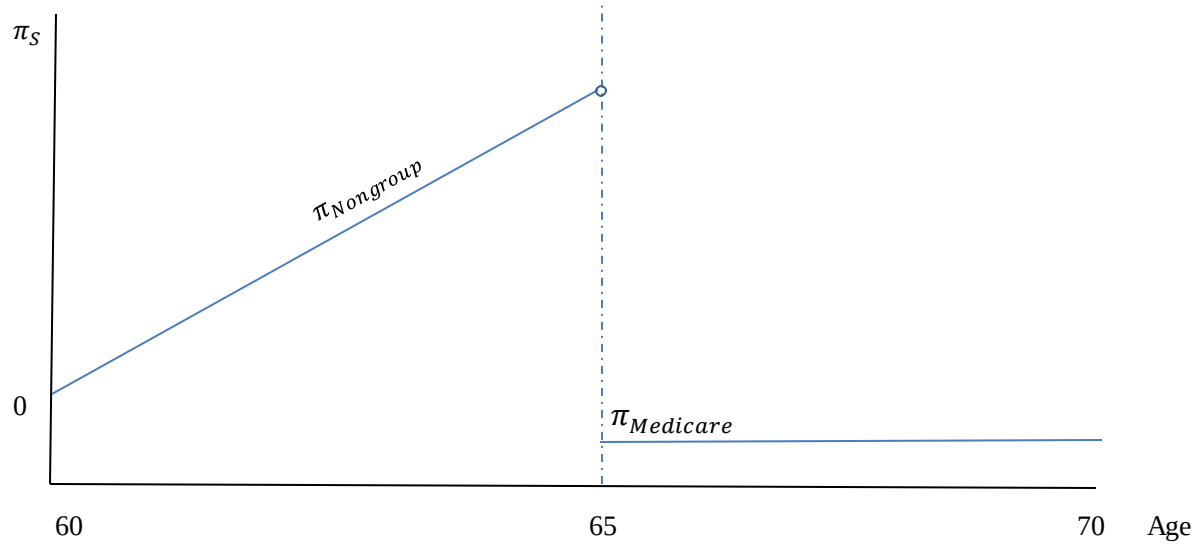


Figure 2: Net Utility Gain from Marriage for Potential Divorcees by Age

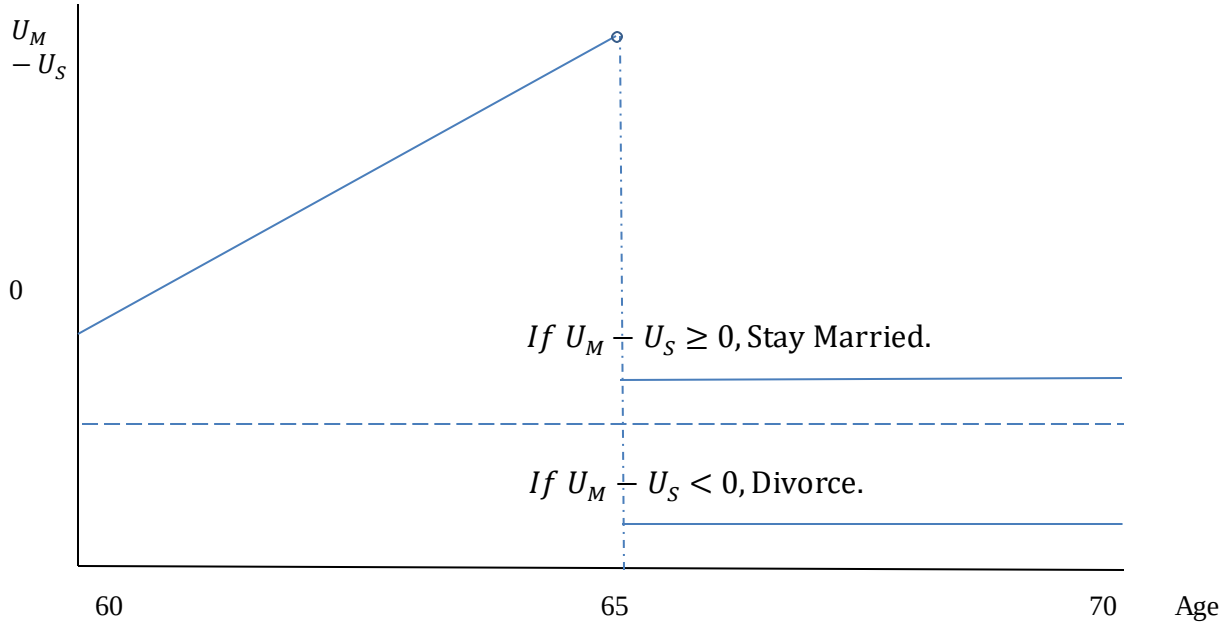
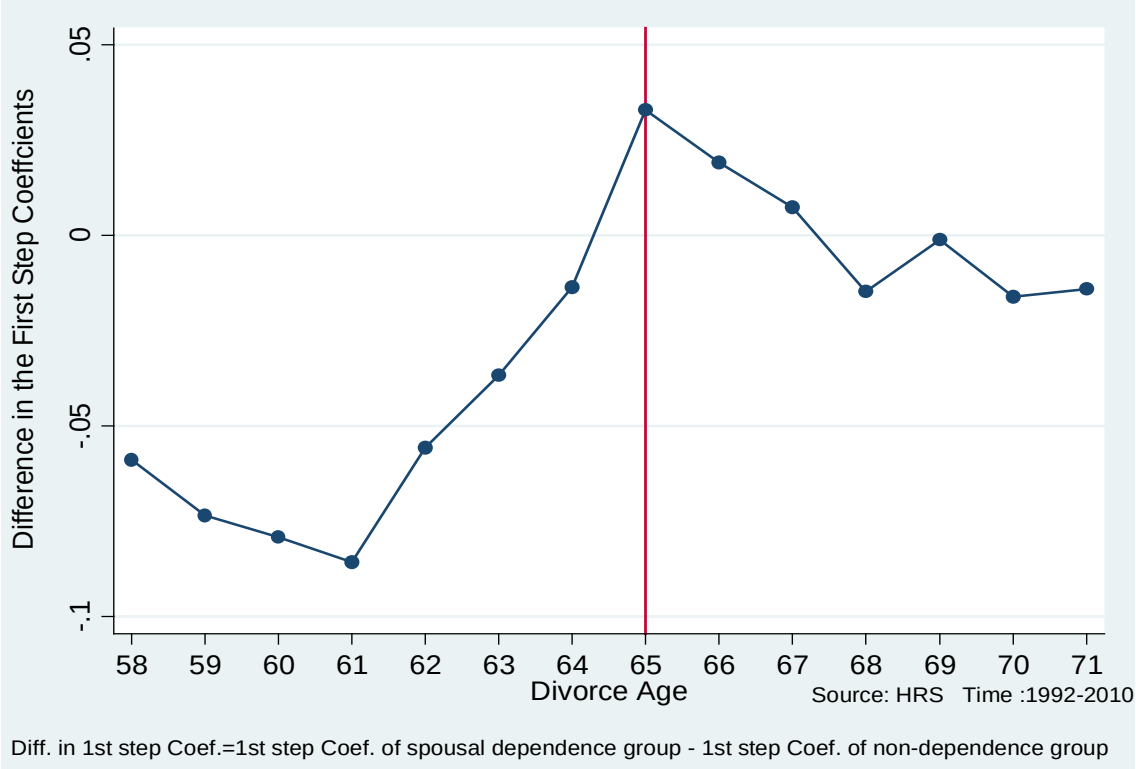


Figure 3: Difference in the First-step Coefficients for the Spousal Coverage Dependent Group and Nondependent Group



Note: Source: HRS 1992–2010. Individuals in the sample are either married or divorced. The spousal coverage dependence group refers to individuals who or whose spouses have a single source of health insurance from the spouse's ESI coverage. "The first-step coefficient for the dependent group" is the estimated coefficient of the interaction terms between the treatment group (ESI Coverage Dependence Group) dummy and age (the older age of the spouses) by using the two-step estimation method. "The first-step coefficient for the nondependent group" is defined similarly for the control group without such ESI coverage dependence. The "difference in the first-step coefficients for the dependent and nondependent groups" is the difference in the first-step coefficient estimates for the two groups.

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Notes

[1] In addition, these probabilities are consistent with the distribution from a small sample with the actual age of divorce reported in the HRS.

[2] The first-step coefficient for the dependent group is the estimated coefficient of the interaction terms between the treatment group (Coverage Dependence Group) dummy and age (the age of the older spouse) using the two-step estimation method. The first-step coefficient for the nondependent group is defined similarly for the control group without such an ESI coverage dependence. The difference in the first-step coefficients for the dependent and nondependent groups refers to the difference in the first-step coefficient estimates for the two groups.

[3] In addition to estimations that are not reported in the attached tables, I exclude retirement in the estimation and obtain similar results for the interaction term between the age cutoff for Medicare and the spousal coverage dependence group to the estimates reported in Tables 1–5 when retirement is included. The results remain robust regardless of whether I control for social security in the estimation.