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‘TIL INSURANCE DO US PART: THE EFFECT OF THE AFFORDABLE CARE ACT PREEXISTING CONDITIONS PROVISION ON MARRIAGE

BY

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‘Til Insurance Do Us Part:

The Effect of the Affordable Care Act Preexisting Conditions Provision on Marriage

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Abstract:

This paper investigates the effect of the 2014 Affordable Care Act preexisting conditions provision on marriage. The policy was implemented to prevent insurers from denying insurance coverage to individuals with health conditions. We test whether the implementation of the provision led to decreases in marriage among affected adults. We add to earlier work on how marital behavior is influenced by policy incentives and examine for the presence of “marriage lock”, a situation in which individuals remain married primarily for insurance. Using data from 2009-2015 and difference-in-differences models, we find that males with preexisting conditions are 5.15 percentage points (6.40 percent) less likely to be married after the policy implementation. Effects are largest for men who had insurance coverage prior to the policy change from a source other than his own employer, suggesting that the inability to attain individual coverage and reliance on spousal insurance provided incentives to remain married.

Keywords: Affordable Care Act; Preexisting Conditions; Health Insurance; Marriage

JEL Classifications: J12, I13, D1

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1. Introduction

In 2010, The Patient Protection and Affordable Care Act (ACA) was enacted by Congress with goals, among others, to ensure that individuals with preexisting conditions are not denied access to affordable health insurance plans. The ACA has been credited with reducing the number of Americans lacking health insurance coverage by about 20 million, leading to a historic low (Avery et al., 2016; Uberoi et al., 2016) and with lessening the effects of job lock among parents of children with chronic conditions (Chatterji et al., 2016). While the reduction of uninsured individuals was a major goal of the ACA, large changes in policy also come with unintended effects. Recent studies have shown that the ACA's dependent coverage mandate, which expanded insurance coverage for young adults by allowing them to remain on parental plans until the age of 26, led to reductions in labor supply (Akosa Antwi et al., 2013; Lenhart and Shrestha, 2016; Colman and Dave, 2018), declines in college enrollment (Jung and Shrestha, 2018), and decreases in fertility (Abramowitz, 2018). It remains important to continue to study the ACA to direct policymakers toward a more developed understanding of the policy.

Our study examines the ACA prohibition of denying individuals insurance coverage based on preexisting conditions, which was implemented on January 1, 2014 for adults, and its impact on marriage.[1] Claxton et al. (2016) estimate that over 27 percent (52 million) of adult Americans under the age of 65 have health conditions that would have made them uninsurable under pre-ACA laws. We test whether the preexisting conditions provision affected marital decisions of individuals who may have been denied coverage prior to enactment of the policy. In a pre-ACA world, it is easy to imagine a scenario in which an individual with health conditions may feel particularly reliant on his or her marriage for reasons of procuring affordable health

[1] The preexisting conditions provision took effect in 2010 for children.

insurance. After implementation of the policy, an individual with preexisting health conditions no longer has to worry about being denied coverage or charged higher premiums by an insurer, and they may feel less reliant on spousal health insurance coverage. This paper contributes to the sparse literature examining how marital behavior is influenced by policy incentives that impact health insurance coverage. Two recent papers have studied the relationship between the ACA young adult provision and marital decisions. Abramowitz (2016) finds evidence that the provision led to a decrease in the likelihood of marriage and an increase in the probability of divorce among young adults. The study provides evidence that while spousal health insurance plans increase the value of marriage, being eligible to remain on parental insurance plans reduces incentives to remain married. In similar work, Barkowski and McLaughlin (2018) highlight the importance of accounting for pre-ACA state-level mandates by showing that young adults who were ineligible for such state-level mandates were less likely to marry prior to the ACA, while increasing their marriage rates following the ACA implementation. While a finding that the ACA dependent coverage provision impacted the marital decisions of young adults is interesting, we believe that an examination of whether the ACA preexisting conditions provision impacted marital decisions of individuals with health issues is also important.

To the best of our knowledge, our study is the first to estimate the impact of the ACA preexisting conditions provision on marital status. Using longitudinal data from 2009 to 2015, and estimating difference-in-differences models, we test whether male household heads with preexisting conditions prior to 2014 changed their marital behavior following the implementation of the policy. We find a 5.15 percentage point reduction in the likelihood of being married for these men compared to those without any preexisting conditions, which corresponds to a 6.40 percent change in marriage rates. We find that the effect is largest for men that were insured prior to 2014, and those that did not have their own employer-sponsored insurance coverage. Additionally, we find a negative and significant effect on marriage among those men who were covered by a spousal, employer-sponsored insurance plan, which we believe to be indication that they may have been particularly reliant on marriage in order to obtain health insurance benefits. Due to the nature of the policy under study, our findings are indication of "marriage lock", a situation in which a couple remains married primarily for one partner to maintain health insurance coverage that he or she could not get otherwise.[2] We show that the results are robust across a variety of specifications including propensity score matching methods and a fully flexible, dynamic difference-in-differences specification introduced by Mora and Reggio (2015). Finally, to alleviate concerns that the effect is being driven by other provisions of the ACA such as dependent coverage and Medicaid expansion, we show that results are largest for those between the ages of 30 and 45 and are not driven by individuals living in Medicaid expansion states.

The rest of the paper is organized as follows. The next section provides background information on the link between health insurance coverage and marriage as well as details regarding the ACA. Section three describes the Panel Study of Income Dynamics (PSID) data, while section four outlines our difference-in-differences, propensity score matching, and Mora and Reggio (2015) methodologies. Section five describes our findings and robustness checks. The final section provides further discussion and concludes.

2. Background

2.1 Marriage and Health Insurance Coverage

[2] Marriage lock is analogous to job lock, i.e., remaining at an employer primarily for insurance purposes.

Gary Becker's economic model of marriage (Becker, 1973, 1974) is based on two basic assumptions: 1) each individual tries to find a partner who maximizes his or her well-being, where well-being is measured by the consumption of household-produced commodities, and 2) the marriage market is in equilibrium. The model shows that gains from marrying compared to remaining single are positively correlated with income, human capital, and relative differences in wage rates. Becker (1973) shows that a necessary condition for marriage is that total output of the marriage must exceed the sum of maximum outputs of single individuals. Spousal health insurance coverage is a component of the value of marriage. A potential result of this could be that unhappily married people may decide to remain in their marriage because they fear that they will not be able to obtain insurance coverage on their own. Abramowitz (2016) notes that the benefit of obtaining insurance coverage through marriage could impact both coupled and single individuals. For single individuals, an available spousal plan may induce individuals to marry since insurance makes marriage more attractive. Similarly, for married individuals, a wider availability of alternative health insurance plans may make marriage less attractive, and induce them to get divorced. We hypothesize that in a pre-ACA world, individuals suffering from preexisting health conditions may be particularly influenced by these marriage channels.

Prior literature has established the link between health insurance coverage and marriage. It has been shown that unmarried women are less likely to be insured than married women (Bernstein et al., 2008; Meyer and Pavalko, 1996), and that divorce leads to losses of coverage among women (Zimmer, 2007; Lavelle and Smock, 2012) and increases in individually-owned private coverage (Peters et al., 2014). Results from a survey in 2008 among a nationally representative random sample of 2,003 adults suggest that insurance might be a determinant of a couple's decision to get married (Kaiser Family Foundation, 2008). Seven percent of survey respondents affirmed that a person in their household got married over the last year mainly to have access to health care benefits. Since individuals with preexisting conditions may have been particularly worried about being denied coverage prior to the ACA, insurance could have played a particularly large role in their marital decisions.

Other work shows that marriage behavior is influenced by policies altering the costs and benefits of marriage. Yelowitz (1998) shows that the expansion of Medicaid eligibility beyond single-parent families significantly increases the likelihood of being married. Similarly, Chen (2013) shows that becoming eligible for Medicare at age 65 is associated with a seven percent increase in the likelihood of getting divorced, suggesting the presence of marriage lock among the near-elderly. Other studies show that marital decisions are influenced by changes to the Aid to Families with Dependent Children program (Moffitt, 1990) and to income taxes (Alm and Whittington, 1997, 1999). Researchers have shown that divorce is not only costly for individuals, but also to society. In a case study conducted for Texas, Schramm et al. (2013) provide evidence that the state spends $3.18 billion on divorce and its related consequences each year, which corresponded to around 12 percent of the state's total budget in 2008. In similar work for Utah, Schramm (2006) estimates that the 9,735 divorces in the state in 2001 cost the state and federal governments nearly $300 million in direct and indirect costs. Both studies emphasize that social policy should strengthen marriages and not provide economic incentives for individuals to get divorced.

2.2 Pre-ACA Medical Underwriting

The ACA preexisting conditions provision was introduced to improve health care access for a vulnerable group of the population: individuals with preexisting health conditions. Prior to the implementation of private insurance market rules in the ACA in 2014, health insurance sold in individual markets was medically underwritten in most states. That is, insurers evaluated the health status, health history, and other risk factors of applicants to determine the type of insurance plans an individual was eligible to receive (Claxton et al., 2016). This resulted in situations where individuals, particularly those with preexisting conditions, faced surcharged premiums, higher deductibles, or in many cases denial of coverage altogether.

At any given time, the vast majority of the population with declinable preexisting conditions could attain health insurance through either an employer or through public programs such as Medicaid, and hence would not be subject to the medical underwriting practices of the individual marketplace. However, in the medium- or long-run, many of these individuals will eventually seek out coverage from the individual insurance market. According to Claxton et al. (2016), the need for individual market insurance is intermittent, and divorce is a common life event that may lead to the disruption of an individual's group coverage and the sudden need for private insurance.

Prior to the ACA, medical underwriting of private insurance coverage posed serious issues for individuals with preexisting health problems. Applications for individual policies often included lengthy questionnaires about the health risk status of applicants and other members of the covered family. Before 2014, these practices were permitted in 45 states as well as in the District of Columbia. Levitt et al. (2013) estimate that, due to medical underwriting, 18 percent of applicants were denied coverage in the individual marketplace prior to the ACA. The authors point out that this estimate is likely a lower bound since it does not include individuals with preexisting conditions who were discouraged from applying for insurance altogether.

3. Data

3.1. Panel Study of Income Dynamics (PSID)

Our study uses data from the Panel Study of Income Dynamics (PSID), a nationally representative longitudinal sample of households and families interviewed annually since 1968 and biannually since 1997. In our analysis, we follow individuals over the years 2009, 2011, 2013, and 2015, which provides us with three survey waves prior to and one survey wave after policy implementation. The PSID gathers information reported by household heads, which by design are male individuals within married-couple families and either male or female individuals within single-parent homes.

Given that the focus of this study is to test whether the preexisting condition provision led to changes in marital status, we examine the effects of the policy change for men since the share of married female household heads who have any preexisting conditions is less than 3 percent of the PSID sample during our survey years. While the main focus of the study is to investigate marital behavior of male household heads, the PSID allows us to test whether behavior is affected differentially depending on whether the head of household or his partner has the preexisting condition. Given that the legal age of marriage in the U.S. is 18, and individuals become eligible for Medicare at the age of 65, our sample consists of individuals between the ages 18 and 64. Additionally, we exclude respondents with missing information on marital status. These restriction leave our analysis with a sample of 10,593 individuals and 42,372 total observations.

In order to narrow our sample to a group of individuals that are likely affected by the preexisting conditions provision, we use responses in the survey to whether a doctor has ever diagnosed the survey participants with any of the following health conditions: stroke, heart attack, heart disease, lung disease, diabetes, cancer, and other serious chronic conditions, which includes conditions such as seizures, kidney disease, autoimmune disorder, Parkinson’s disease, coronary problems, and bone disorder. The main treatment group of our study consists of individuals who report that they had at least one of the conditions in all three pre-policy survey years. This provides us with a treatment group of 1,534 individuals, for which we have 6,136 total observations. For the control group, we consider people who report that they have no health conditions in all years prior to the reform (9,059 individuals and 36,236 observations).[3]

The PSID provides several advantages to study the effect of the policy on marital behavior including its longitudinal nature, the fact that it covers individuals from all age groups and collects information on a number of health conditions. However, there are still some limitations with using the PSID. While we are using the latest available survey year (2015), this means that our analysis is only able to capture the early effects of the policy. Given that divorces can be rather lengthy and take significant time to process, we believe the results could be larger when including observations several years after the policy implementation. Another limitation of the PSID is that it only interviews household heads. This clearly limits us to focusing on the effects on marital behavior within a sample male household heads. Finally, an ideal data set for this study would provide information on whether respondents had actually been denied insurance coverage in the past, which is not available in the PSID or in other data sets as far as we are aware.

3.2 Summary Statistics

Table 1 provides summary statistics for the treatment and control group of the analysis. The statistics indicate that men forming the treatment group are 4.80 percentage points less likely to be married after policy implementation compared to the pre-policy survey years. Conversely,

[3] In an alternative specification, we find qualitatively similar results using a treatment group of individuals who report having had at least one of the conditions in 2013 (3,503 individuals).

the likelihood of being married increases for members of the control group by 1.47 percentage points. The decline in the share of men with preexisting conditions who are married after the policy change is confirmed by Figure 1, which provides graphical evidence of changes in marital status between 2009 and 2015 for individuals in our sample. While relatively similar trends in the likelihood of being married are noticeable for the two groups during the years prior to the policy change, Figure 1 shows a larger decline in the likelihood of being married for individuals with preexisting conditions between 2013 and 2015. In order to check if individuals were potentially delaying divorce until after 2014, we examined aggregate U.S. divorce rates in the years prior to the policy change. According to statistics from the Center for Disease Control (CDC) and the National Center for Health Statistics (NCHS), divorce rates per 1,000 total population remained fairly similar during the pre-ACA years and even slightly declined between 2011 and 2016 (CDC/NCHS, 2017).

The summary statistics in Table 1 show that treated household heads are on average older and less likely to work compared to those in the control group, while the two groups are relatively similar in terms of race, education, and total family income. The bottom of Table 1 shows statistics for the share of individuals in the treatment group reporting to suffer from various health conditions in the PSID during the pre-treatment period (2009, 2011 and 2013). The most prevalent condition is diabetes (44.59 percent), while shares of individuals who have had a stroke, heart attack, cancer, or heart and lung diseases are substantially smaller.

4. Econometric Methods

4.1 Difference-in-Differences

Our study exploits the ACA prohibition of denial based on preexisting conditions provision to test for the effects of the policy change on marriage behavior. In the main empirical specification, we employ a difference-in-differences (DD) framework to observe the average treatment effects on a group of individuals that is most likely affected by the policy implementation. The main treatment group consists of heads of household who report that they have been diagnosed with at least one serious health condition in all survey observations prior to the policy change (2009, 2011, and 2013). In an alternative specification, individuals whose spouse had a preexisting condition throughout the pre-policy period form the treatment group. The control group consists of individuals who had no preexisting health conditions between 2009 and 2013. We estimate the following equation to obtain average treatment effects:

$$Y_{ist} = \beta_0 + \beta_1 \, Treat_{ist} + \beta_2 \, X_{ist} + \delta_{DD} \, Post_{ist} * Treat_{ist} + \lambda_1 \, Year_t + \lambda_2 \, State_s + \alpha_i + \varepsilon_{ist}, \qquad (1)$$

where Y_{ist} is an indicator that equals one if individual i living in state s at time t is married and zero if he is either unmarried, widowed, divorced or separated. $Treat_{ist}$ equals one if the individual suffers from a preexisting condition, while $Post_{ist}$ equals one in the post-treatment period (2015) and zero in the three pre-treatment years (2009, 2011, and 2013). The inclusion of the vector X_{ist} controls for observable individual characteristics, such as age, the number of children in the household, education, employment status, and total household income.[4]

The main parameter of interest is δ_{DD}, which captures the effect of the policy on the likelihood of being married. Equation (1) controls for year and state fixed effects to account for existing differences in marital status across time and space. One concern is that if unobservable characteristics at the state-level cause divorce to spike in certain states, and individuals making up the treatment group may disproportionately live in these states. In additional models, we address this concern by including state-specific time trends. Finally, equation (1) also includes

[4] Additionally, we re-estimate our model when excluding employment status and total household income and find that the results remain almost unchanged.

individual fixed effects (α_i), which allows us to account for time-invariant heterogeneity across individuals. The use of longitudinal data reduces any potential bias due to changes in the composition of the sample before and after the policy change. We estimate each of our DD specifications using linear probability models, with the standard errors clustered at the state level.

While the main DD estimation for this study uses all heads of household who had preexisting conditions throughout the pre-policy period as the treatment group, we further exploit the longitudinal nature of the PSID in two additional DD specifications. First, we narrow the sample to individuals who were married in the pre-treatment years (2009-2013). Again, the two treatment groups consist of households where either the head or the spouse has at least one health condition before the policy change, respectively. This specification allows us to examine the effects of the policy on the flow of marriages more directly rather than the marriage stock. Using the stock of married people as an outcome could underestimate the magnitude of the effects of the provision (Abramowitz, 2016), and observing changes in marriage may capture policy effects more precisely.[5]

Second, we use information on insurance status and narrow our sample even further to household heads with preexisting conditions who are covered by their spouse's insurance plan in the pre-policy years. While the sample size is smaller, this specification estimates treatment effects on a group of individuals that is most likely affected by the policy, since they may be particularly reliant on their marriage for health insurance. Thus, the results from this specification are closer to treatment effects on the treated.

[5] Similarly, Abramowitz and Dillender (2017) argue that using stock outcomes and flow outcomes may yield disparate results.

4.2 Alternative DD Specifications

4.2.1 Propensity Score Matching Difference-in-Differences Model

To provide further robustness to the validity of our main DD estimates, we estimate propensity score matching DD models. This method allows us to compare the distribution of outcomes between individuals in the treatment group and their matched counterparts in the control group without having to make any functional form assumptions. Thus, our estimated treatment effects are weighted averages of the difference-in-differences between each of the treated individuals and his matched control.

We use estimated propensity scores, which calculate the probability of treatment given a vector of observable characteristics, to match individuals who are treated to those who are similar but are not impacted by the policy change. The propensity scores are based on pre-treatment variables and are estimated using probit models. Observable characteristics that are included to obtain the propensity scores are age, race, education, number of children, and state of residence. Following Rosenbaum and Rubin (1983), the use of a function of X, called the propensity scores P(X), rather than a potentially high-dimensional vector of covariates implies that:

$$E\,(Y_0 \mid D = 1, P(X)) = E\,(Y_0 \mid D = 0, P(X)), \tag{2}$$

where Y_0 denotes the untreated state, D = 1 indicates treatment, and D = 0 indicates non-treatment. Our analysis follows Heckman et al. (1998) DD matching methods, which use both comparisons between treated and non-treated, and differencing over time. Thus, the conditions needed to identify the average treatment effect on the treated (ATET) using the DD matching estimator are:

$$E\ (Y_{0,t} - Y_{0,t'} \mid D = 1, P(X)) = E\ (Y_{0,t} - Y_{0,t'} \mid D = 0, P(X)), \qquad (3)$$

where t and t’ represent the post- and pre-treatment periods, respectively. Thus, the average treatment effect (ATE) provides a weighted average of the difference-in-differences between individuals in the treatment and control groups, and it is obtained by estimating the following equation:

$$ATE_{DD} = E(Y_{1t} \mid D=1, P(X)) - E(Y_{0t} \mid D= , P(X)) - E(Y_{1t'} \mid D=1, P(X)) - E(Y_{0t'} \mid D=0, P(X)). \quad (4)$$

Our empirical analysis uses both nearest neighbor and kernel density matching on the propensity scores (Becker and Ichino, 2002). Standard errors are obtained following Abadie and Imbens (2016), who established how to account for the fact that propensity scores are estimated in the first stage. Figure 2 shows a balance plot for the propensity score matching conducted in our analysis. The figure suggests that our matching is successful and allows comparing the effects of the policy change between similar individuals in each the treatment and control groups.

4.2.2. Mora & Reggio DD Model

DD models require an assumption that trends in the variable of interest are similar for both treatment and control groups in the absence of the policy change. This assumption implies that without the treatment, differences between the groups are assumed to be time-invariant. Mora and Reggio (2015) point out that the identification of the treatment effect does not only depend on the parallel trends assumption, but also on the trend modeling strategy applied by researchers. For example, Mora and Reggio (2015) show that DD estimates will differ substantially depending on whether group-specific linear trends or group-specific, time-invariant linear trends are included in the analysis in order to accommodate for trend differentials between treatment and control groups. By arguing that researchers often overlook this fact, the authors introduce an alternative DD estimator, which identifies the effect of the policy using a fully flexible dynamic specification and includes a family of alternative parallel growth assumptions (Mora and Reggio, 2015). The two main advantages the authors list in favor of their DD estimate are that it: 1) allows for flexible dynamics and for testing restrictions on these dynamics; 2) does not impose equivalence between alternative parallel assumptions. Mora and Reggio (2015) show that this alternative estimator is acquired in two steps. In the first step, standard least squares estimation of the fully flexible model is conducted. In the second step, the solution of the equation in differences identifies the treatment effects. The Mora and Reggio (2015) fully flexible DD estimator is obtained from the following equation:

$$E\left[Y_t \,|D\right] = \delta + \sum_{\tau=t_2}^{T} \delta_\tau I_{\tau,t} + \gamma^D D + \sum_{\tau=t_2}^{T} \gamma_\tau^D \times I_{\tau,t} \times D$$

where $I_{\tau,t}$ is a dummy for period τ and γ^D is a control for group differences in linear trends. Estimating this alternative model allows us to test the validity of standard DD assumptions made in our baseline specification, and a finding that the estimates obtained from this model are consistent with baseline estimates can provide evidence for additional robustness of the main results.

4.3 Additional Models

In addition to estimating alternative DD models, we conduct several robustness checks. We run two additional specifications that account for the ACA Medicaid expansions that took place in 26 states and DC in 2014, the same year as the preexisting conditions provision.[6] We

[6] The states that expanded Medicaid in 2014 are: Arizona, Arkansas, California, Colorado, Connecticut, Delaware, District of Columbia, Hawaii, Illinois, Iowa, Kentucky, Maryland, Massachusetts, Michigan, Minnesota, Nevada, New Hampshire, New Jersey, New Mexico, New York, North Dakota, Ohio, Oregon, Rhode Island, Vermont, Washington, and West Virginia.

separately examine the impact of our policy change on marriage behavior of individuals with preexisting conditions in states that expanded Medicaid and the 24 states that did not. If Medicaid expansion is driving any differences in marriage rates between individuals with and without preexisting conditions, we would expect no changes in marital status in states that did not expect Medicaid.

Another ACA provision that could have affected marital behavior is the 2010 dependent coverage mandate, which allows young adults to remain on parental insurance plans until the age of 26. Abramowitz (2016) provides evidence that the mandate reduced marriage rates among young adults. While the policy examined in this study was implemented four years later, any observed effects might still be related to the dependent coverage mandate. To empirically test for whether the dependent coverage mandate is driving any observed marriage effects, we re-estimate our main DD model for four different age groups. If the 2010 dependent coverage provision drives our results, we would expect to find larger effects for individuals below the age of 27.

We also check whether the effects differ between people who were insured and uninsured before the policy change. If an individual was uninsured prior to the policy, then we would expect to see no significant effect of the preexisting condition provision on their marital decisions (as they would not have been reliant on spousal insurance coverage in the pre-policy period). We also check for differences between those who had their own employer-sponsored coverage and those who did not. As the policy is most likely to impact individuals that are unable to attain their own employer-sponsored insurance coverage, we expect to see larger effects for those without this form of coverage in the pre-policy period.

To provide additional validity to our main findings, we estimate a falsification and three placebo tests. A potential concern is that individuals with preexisting conditions are quite different than those without, and perhaps it is the worsening of one's condition that eventually leads to a divorce. To mitigate concerns that differences in health outcomes might be driving any changes in marriage behavior, we estimate another DD specification where individuals with no preexisting conditions but poor self-reported health form the treatment group and those who have no conditions and are not in poor health form the control group. Since individuals making up the treatment group do not suffer from any preexisting conditions that would make them uninsurable under pre-ACA law, we would expect to see no significant effects of the preexisting conditions provision on their marital decisions. Finally, in three placebo tests, we re-estimate our main analysis for the time periods 2003-2009, 2005-2011 and 2007-2013 and create artificial treatment indicators. None of these time periods had changes in policy directly targeting the insurance needs of individuals with health conditions. Thus, if estimates using these artificial treatment indicators are statistically significant, it may be indication that our main results are spuriously driven by other changes that differentially affected individuals with and without preexisting health issues.

Results

5.1 Main DD Model

Table 2 presents the main DD estimates from our analysis. In the main specification in Panel A, which includes all fixed effects and controls, we find that the policy change reduced the likelihood of being married among male household heads with preexisting conditions by 5.15 percentage points ($p<0.01$). This corresponds to a 6.4 percent reduction compared to the baseline mean (0.805). We furthermore find that the results remain unchanged when adding state-specific time trends.[7] In additional specifications, shown in Appendix Table A1, we include additional controls and show that the results are not driven by marriage or health history. We find that the main results remain almost unchanged when including controls for marriage length, the number of marriages, as well as the age when the preexisting condition first occurred.[8] Panel A of Table 2 also shows that the policy reduced the likelihood of being married by 3.24 percentage points for household heads whose spouse had a preexisting condition in the survey years prior to 2014 ($p<0.05$). This suggests that marriage is affected independent of which partner has the preexisting condition.

Panel B narrows down the sample to individuals who were married throughout the pre-policy period in order to test whether the flow of marriages was impacted by the reform. Again, we find negative and statistically significant effects on the likelihood of remaining married for heads of household with health conditions as well as for individuals whose spouses have preexisting conditions. Finding consistent results for both the stock and flow of marriages provides additional evidence that the policy change affected marital behavior.

Finally, we narrow the sample in Panel C to household heads who were covered by their spouse's insurance plan in the years prior to 2014. As they would likely have been highly dependent on the marriage to attain affordable health insurance, individuals with preexisting conditions within this subsample are those who are most directly affected by the reform. In the full specification, and despite a low number of observations and potential issues with statistical power (325 treated individuals), we find a 3.93 percentage point decline ($p<0.05$) in the

[7] Six states (Maine, Massachusetts, New Jersey, New York, Vermont, and Washington) did not request detailed information about a person's medical history when they tried to purchase insurance prior to the ACA. We estimated our main model excluding these states and found a decline in the likelihood of being married of 5.13 percentage points ($p<0.01$), which is almost identical to the main result.

[8] In additional specifications, we exclude people in the control group whose spouse have at least one preexisting condition at some point during our study. The results remain unchanged are available upon request.

likelihood of being married for this group of individuals. This furthermore suggests that the ability to obtain insurance coverage independently of one's partner influenced people's decisions of remaining married.

5.2 Additional Results

Table 3 presents estimates from several models testing the robustness and validity of our findings. Panel A shows treatment effects obtained from two different propensity score matching DD specifications. While the magnitudes for the nearest neighbor and kernel matching differ, they both provide evidence for of a statistically significant negative effect of the policy on the likelihood of being married (both $p<0.01$). Given that these estimates are in line with the standard DD estimates removes some concern about observable differences in the treatment and control groups (Table 1) and that these differences are driving the observed changes in marital behavior. Similarly, Panel B shows a 4.03 percentage point reduction ($p<0.05$) in marriage when estimating the alternative DD model proposed by Mora and Reggio (2015), which provides evidence that the main DD estimates are robust to a more flexible, dynamic specification and to different parallel trends assumptions for the two groups.

Panel C shows estimates for three placebo tests using different time periods and artificial policy indicators to ascertain that our main findings are not spuriously driven by other events or policy changes. The results from these models show that the artificial policy indicators have no effects on marriage behavior of individuals with preexisting conditions, which further suggests that the 2014 change in policy is driving our main findings.

While the treatment group in our main analysis is formed by individuals who have at least one health condition in all three pre-policy survey years, we present estimates for each preexisting condition in Panel D. With the exception of heart attack, which shows an effect close to zero in magnitude, we find negative effects on marital behavior among all health conditions. It should be noted, however, that only the estimates for diabetes and other chronic conditions are statistically significant (both $p<0.01$). One explanation for this could be that, as shown in Table 1, the share of individuals reporting these two conditions in substantially higher than for all other conditions.[9] Thus, the lack of significance for the other health conditions may be related to a lack of statistical power.

Panel E shows that the prohibition denying insurance based on preexisting conditions reduced the number of children in the household by 0.158 ($p<0.05$). According to statistics provided by the U.S. Census Bureau, mothers receive custody after divorces in more than 80 percent of cases (U.S. Census Bureau, 2015). Thus, the observed decline in the number of children living with male household heads with preexisting condition is in line with the decrease in marriage rates among this group.[10]

Finally, we estimate a falsification test for which we assign individuals without any preexisting conditions but poor self-reported health in all pre-policy survey waves to the treatment group, while individuals with no poor health and no health conditions form the control group. Consistent with the fact that these “treated” individuals should not be affected by the ACA preexisting conditions provision, we do not find a statistically significant marriage effect for this alternative specification.

5.3 Heterogeneous Effects

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[9] When further separating the different conditions groups under “other chronic conditions”, we find that marriage behavior was most affected for people suffering from bone/joint disorders as well as autoimmune disorders or lupus. The results for these additional “other chronic conditions are available upon request.

[10] The fact that the number of children is affected by the policy change suggests that it might me endogenous in the main specification examining the effects on marriage behavior. We find that our main results are unchanged when excluding the number of children from the model.

Table 4 provides estimates for several subgroups of the population. Panel A presents results for individuals by pre-policy insurance status. The results suggest that our main findings are driven by individuals who had insurance coverage in the survey years prior to 2014. It appears unlikely that individuals that were uninsured prior to the policy would have felt tied to marriage due to health insurance. This result is in line with our previous finding of a significant effect among individuals that were covered by their spouse's plan prior to the reform (Table 2, Panel C). In addition, we show that the effects are mainly driven by individuals who did not have their own employer-sponsored coverage before 2014. This would likely be a subgroup that is most likely to be reliant on spousal coverage. The estimates in Panel A provides further robustness to our main results since these individuals were likely to be most affected by the provision.

Next, we examine whether the 2014 Medicaid expansions and the 2010 dependent coverage mandate, which were both part of the ACA, are driving the finding of our study. When separately testing for the effect of the preexisting conditions provision on marriage behavior in states that expanded Medicaid in 2014 and those that did not, we find statistically significant reductions in the likelihood of being married for both groups of states (Panel B). While the effect is larger in magnitude for states that experienced an expansion, a significant finding in non-expansion states suggests that Medicaid changes are not driving the main results of this study. Panel C provides treatment effects for four age groups. While our results suggest that marriage behavior is affected across all groups (all $p<0.01$), the largest effect is found for individuals between ages 30 and 45, and the smallest effect is found for young adults between 18 and 26. While Abramowitz (2016) provides convincing evidence that the 2010 dependent coverage mandate affects marital decisions of young adults, Panel C provides additional evidence that the results in our study are the result of the 2014 preexisting provisions provision and are not driven by the dependent coverage provision.

Panel D shows the effects of the policy also differed across education levels. While marriage behavior of individuals with less than 12 years of completed education is not significantly affected, we find reductions in the likelihood of being married of 4.88 percentage points ($p<0.10$) and 6.36 percentage points ($p<0.01$) for heads of household with 12 and more than 12 years of education, respectively. It seems plausible that these heterogeneous effects across education levels are due to varying levels in the understanding of health insurance and changes in policy, i.e., educated individuals are more likely to be aware of the preexisting conditions provision and its role in obtaining insurance coverage, and thus are more likely to respond to the policy. Also, highly educated people might be more likely to afford their own health insurance upon being eligible to do so after the reform.

6. Conclusion

Our findings indicate that the ACA preexisting conditions provision increased marital dissolution among men with health issues. Our main estimate indicates that male household heads with preexisting conditions are 6.40 percent less likely to remain married after the change in policy. Additionally, we find negative effects for males with preexisting conditions covered by spousal employer-sponsored insurance, a subgroup that would have been particularly reliant on a marriage to maintain coverage in the pre-ACA period. We show that our estimates are robust across a variety of specifications, including those that utilize propensity score matching. To mitigate concerns that the findings are driven by other parts of the ACA, such as the dependent coverage provision or Medicaid expansion, we conduct subgroup analyses to show that effects are not driven by young adults and are not unique to those living in Medicaid expansion states.

Our estimates provide evidence that individuals with health conditions may have been particularly reliant on marriage prior to the ACA since insurance companies were able to deny them insurance coverage. We believe that our results are indication of marriage lock, i.e., remaining in a marriage primarily for health insurance purposes. The magnitudes of our findings are comparable with previous estimates of the effects of insurance access and changes in policy on marriage behavior. Abramowitz (2016) finds that the dependent coverage mandate decreased marriage rates by 8.8 to 9.3 percent among young adults, while Chen (2013) estimates a 7 percent increase in the number of divorces for individuals with spousal insurance coverage upon achieving Medicare eligibility at age 65.

By preventing insurers from denying coverage to individuals with preexisting conditions, the ACA may have added a level of relationship flexibility for people with health conditions. By freeing individuals from concerns that they may be denied coverage elsewhere or charged higher premiums, the preexisting conditions provision presumably lowers the benefits of marriage by making spousal insurance coverage less valuable. While likely not having an impact on the average person, the policy has the potential to influence marital decisions of those with preexisting conditions. Given the pre-ACA laws, individuals with preexisting conditions are most susceptible to marriage lock since the value of a spousal health insurance plan is amplified for those that have fears of being denied coverage elsewhere. Marriage lock among individuals with chronic conditions is a topic that is under-studied, and this paper both offers a better understanding of the insurance struggles of people with preexisting conditions and provides new evidence on the effects of the ACA on relationship dissolution outcomes. It is important to continue to track trends in marital behavior and other long-term outcomes related to well-being of individuals with preexisting conditions. Researchers should continue to examine potentially unintended consequences of the ACA and similar changes in policy to better understand the welfare implications of such reforms.

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Figure 1: Proportion Married Across Time

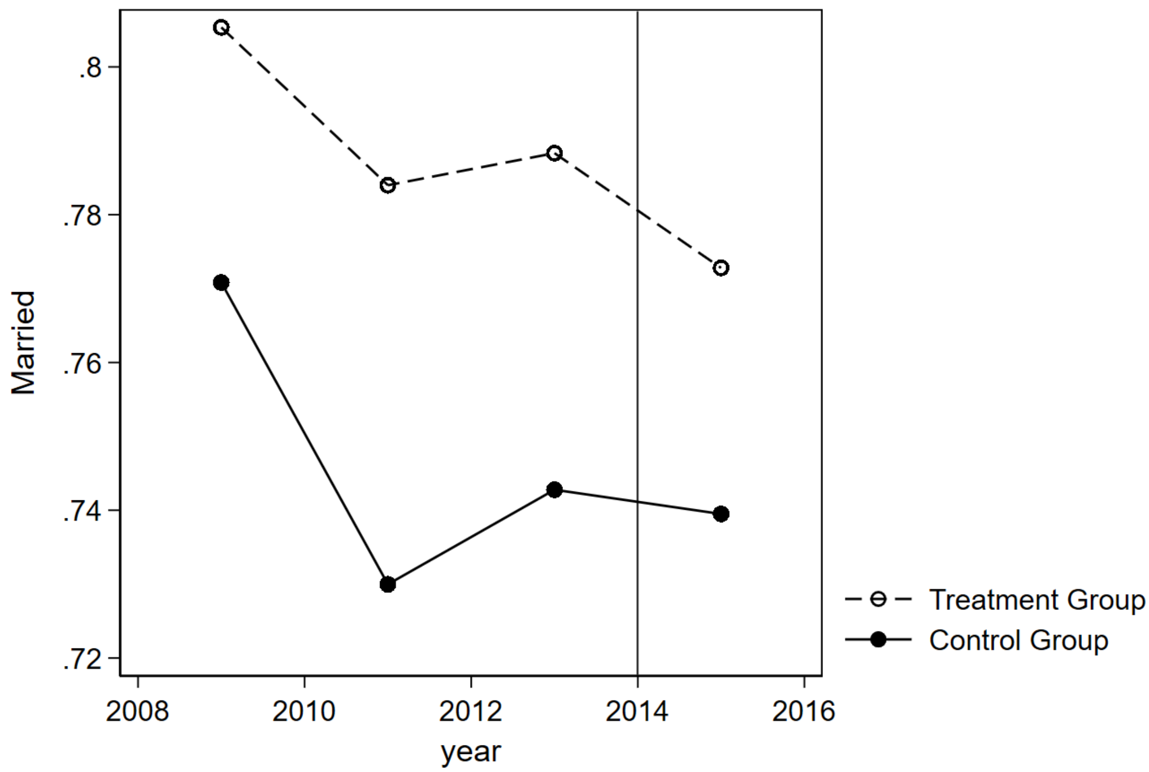

Figure 2: Propensity Score Matching Balance Plot

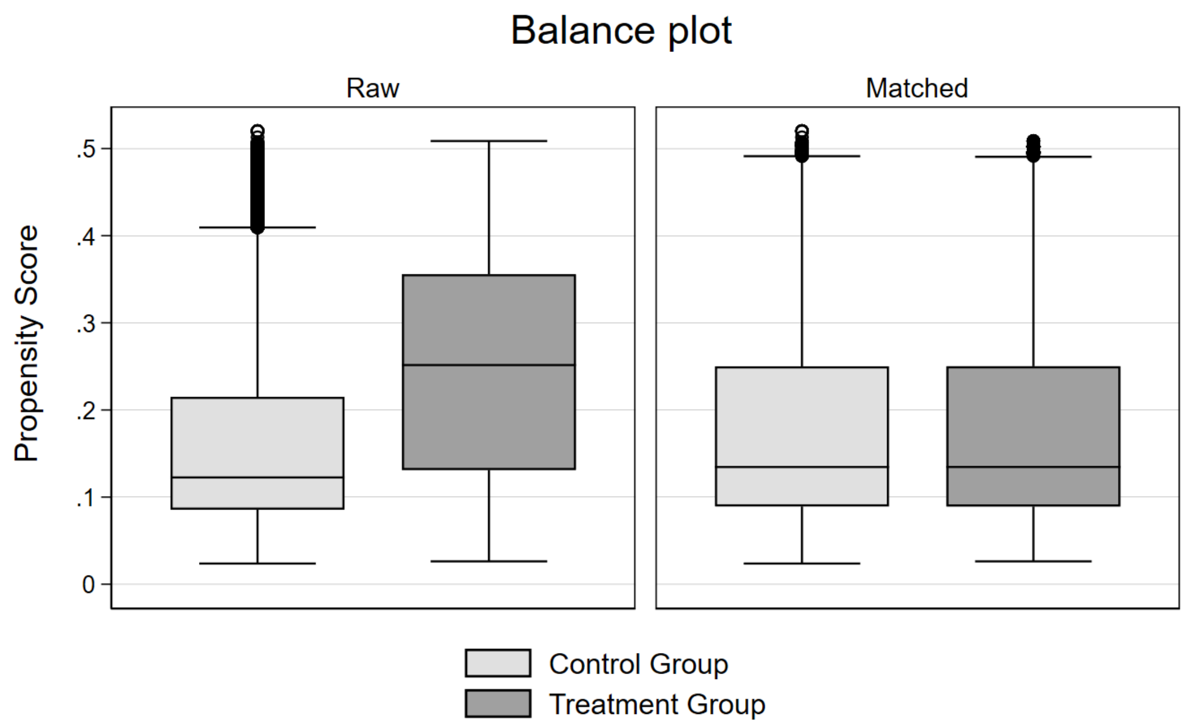

Table 1: Descriptive Statistics

	Treatment Group	Control Group
Married		
Pre	0.8051	0.7474
	(0.3961)	(0.4345)
Post	0.7571	0.7621
	(0.4290)	(0.4258)
Age	47.7825	39.0415
	(10.5326)	(10.3855)
Working	0.6598	0.8638
	(0.4738)	(0.3431)
# Children in HH	1.0495	1.4942
	(1.2727)	(1.4018)
White	0.6588	0.6361
	(0.4741)	(0.4811)
Black	0.2836	0.3001
	(0.4508)	(0.4583)
Other race	0.0575	0.0638
	(0.2329)	(0.2445)
<12 years of education	0.1488	0.1500
	(0.3559)	(0.3571)
12 years of education	0.3015	0.2915
	(0.4589)	(0.4545)
>12 years of education	0.5497	0.5585
	(0.4976)	(0.4966)
Total Household Income	$83,936.40	$84,624.29
	(66,654.37)	(62,092.47)
Stroke (pre)	0.0831	-
	(0.2762)	
Heart attack (pre)	0.1504	-
	(0.3575)	
Heart disease (pre)	0.1430	-
	(0.3502)	
Lung disease (pre)	0.1117	-
	(0.3150)	
Diabetes (pre)	0.4459	-
	(0.4971)	
Cancer (pre)	0.1548	-
	(0.3618)	
Other serious condition (pre)	0.3503	-
	(0.4771)	
N	6,136	36,236

Table 2: The Effects of the Policy Change on Marital Status

Panel A: All male heads of households	**Likelihood of being married**		
Has a condition	-0.0603***	-0.0515***	-0.0513***
	(0.0163)	(0.0161)	(0.0162)
Sample Mean		0.8051	
Spouse has a condition	-0.0530***	-0.0324**	-0.0301**
	(0.0135)	(0.0151)	(0.0153)
Sample Mean		0.9280	
Panel B: Male heads of HH married between 2009 and 2013			
Has a condition	-0.0299**	-0.0306**	-0.0298**
	(0.0133)	(0.0134)	(0.0134)
Spouse has a condition	-0.0359***	-0.0345***	-0.0355***
	(0.0124)	(0.0120)	(0.0120)
Panel C: Male heads of HH covered by wife's ESI pre-policy change			
Has a condition	-0.0473**	-0.0393**	-0.0404**
	(0.0185)	(0.0175)	(0.0184)
Fixed effects	x	x	x
Control variables		x	x
State-specific time trends			x

Robust standard errors, clustered by states, are shown in parentheses. The control variables include age, race, education, employment status and household income. * p< 0.10, ** p< 0.05, *** p < 0.01.

Table 3: Additional DD Results

	DD Effect	N
Panel A: DD Propensity Score Matching		
Nearest Neighbour	-0.0437***	41,663
	(0.0031)	
Kernel Matching	-0.1356***	41,663
	(0.0059)	
Panel B:		
Alternative DD model (Mora and Reggio)	-0.0403***	42,372
	(0.0154)	
Panel C: Placebo Tests		
Pre: 2003-2005, Post 2007-2009	0.0056	40,432
	(0.0040)	
Pre: 2005-2007, Post 2009-2011	-0.0011	42,391
	(0.0031)	
Pre: 2007-2009, Post 2011-2013	0.0009	44,822
	(0.0034)	
Panel D: Type of Condition		
Stroke	-0.0597	31,009
	(0.0377)	
Heart attack	-0.0011	31,326
	(0.0334)	
Lung disease	-0.0124	31,276
	(0.0266)	
Diabetes	-0.0722***	33,066
	(0.0183)	
Cancer	-0.0203	31,367
	(0.0230)	
Other chronic condition	-0.0391***	32,648
	(0.0116)	
Panel E:		
Effects on # of children in HH	-0.1583**	42,372
	(0.0670)	
Panel F: Falsification Test		
Poor health status as treatment indicator	-0.0216	36,133
	(0.0586)	

Robust standard errors, clustered by state, are shown in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 4: Heterogeneous DD Results

	DD Effect	Sample Mean	N
Panel A: Pre-2014 Insurance Status			
Uninsured	-0.0197	0.6690	7,730
	(0.0151)		
Insured	-0.0454***	0.8302	33,933
	(0.0038)		
Employer-Sponsored Coverage	-0.0181	0.9154	18,714
	(0.0140)		
No Employer-Sponsored Coverage	-0.0809***	0.7175	23,658
	(0.0220)		
Panel B: ACA Medicaid Expansions			
Expansion States	-0.0657***	0.8278	20,788
	(0.0234)		
Non-Expansion States	-0.0456**	0.7822	21,255
	(0.0228)		
Panel C: Age			
18-26	-0.0338***	0.4270	4,002
	(0.0118)		
27-29	-0.0547***	0.7412	3,621
	(0.0114)		
30-45	-0.0961***	0.7694	20,200
	(0.0059)		
46-64	-0.0430***	0.8386	13,840
	(0.0051)		
Panel D: Education			
< 12 Years	0.0262	0.6935	6,242
	(0.0400)		
12 Years	-0.0488*	0.7438	12,207
	(0.0273)		
> 12 Years	-0.0636***	0.8780	23,214
	(0.0217)		

Robust standard errors, clustered by state, are shown in parentheses. * p< 0.10, ** p< 0.05, *** p < 0.01.

Table A1: The Effects of the Policy Change on Marital Status
(with Marriage and Health History Controls)

	Likelihood of being married			N
Panel A: Controls for marriage history				
DD Effect	-0.0463***	-0.0382***	-0.0344**	42,372
	(0.0154)	(0.0141)	(0.0141)	
Sample mean		0.8051		
Marriage length	x		x	
Number of marriages		x	x	
Panel B: Controls for health history				
DD Effect	-0.0507***	-0.0525***	-0.0534***	42,372
	(0.0160)	(0.0158)	(0.0159)	
Sample mean		0.8051		
Age condition first occurred - head	x		x	
Age condition first occurred - spouse		x	x	

Robust standard errors, clustered by states, are shown in parentheses. All models control for employment status, the number of children in the household, race, and education. Furthermore, individual, state and year fixed effects are controlled for. * p< 0.10, ** p< 0.05, *** p < 0.01.