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Tom Potoms
Sarah Rosenberg

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Public insurance and marital outcomes: Evidence from the Affordable Care Act's Medicaid expansions

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Public insurance and marital outcomes: Evidence from the Affordable Care Act's Medicaid expansions*

Tom Potoms† Sarah Rosenberg‡

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Abstract

Public insurance may disincentivize marriage, but for whom? Using the American Community Survey, we find that an increased likelihood of Medicaid eligibility owing to the Affordable Care Act reduces marriage rates, particularly among people with higher education levels. We develop a search model of the marriage market that shows that those with a high expected financial surplus from marriage respond more to public insurance, as it provides an outside option against financial risk. These findings suggest a new hypothesis for the marriage gap across the socioeconomic spectrum: a larger material surplus from marriage increases willingness to marry, even with lower match quality.

Keywords: Marital quality, Health insurance, Medicaid expansions, Marriage market

JEL: D10, I38, J12

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†University of Sussex. Jubilee Building, Falmer, Brighton, BN1 9SL, UK.

‡Lund University, School of Economics and Management, Box 7080, SE-220 07 Lund Sweden. E-mail:sarah.rosenberg@nek.lu.se

1 Introduction

Marriage provides an important source of insurance against aggregate and individual shocks. Most high-income countries provide public insurance to individuals, and there is a long history of debate regarding whether such programs incentivize divorce and disincentivize marriage by making singlehood more financially attractive. This debate has particularly focused on people with low socioeconomic status, a group with lower marriage rates over recent decades than those with higher socioeconomic status. This marriage gap is often raised in discussions of intergenerational mobility and inequality (Kearney, 2023).

We hypothesize that public insurance may affect marriage for a broader range of people than has previously been recognized. Although there is widespread recognition that the financial value of singlehood can affect the decision to enter marriage, there is less awareness that the financial value of marriage might also have an impact. Importantly, the financial value of marriage differs markedly by socioeconomic status: The value of legal marriage as compared to remaining single or cohabitating is lower for individuals with less income and wealth. This feature owes in large part to positive assortative matching on income and education, where individuals with less income and wealth are more likely to partner with others of similar socioeconomic status.

If there is little or no material value for marriage to begin with, an improvement in the outside option is unlikely to change marriage decisions. By contrast, if there are large material gains to marriage, a change in the financial value of singlehood may change decisions about marriage. In this context, it is likely that individuals who would receive higher material gains from marriage will be more sensitive to any crowding-out effect from increases in (often means-tested) public insurance programs. Likewise, an expansion of public insurance programs will not necessarily disincentivize marriage for lower income individuals if marriage does not appear to yield a financial benefit.

Research on such effects of public insurance across the socioeconomic spectrum are, however, rather scarce. There is a substantial body of work that has established that changes in insurance programs and regulations do affect marital decisions overall, whether it is through public health insurance expansions (Yelowitz, 1998; Decker, 2000; Farley, 2001; Slusky and Ginther, 2021; Hampton and Lenhart, 2021; Chatterjee, 2024), private health insurance mandates (Abramowitz, 2016; Heim et al., 2018; Barkowski and McLaughlin, 2022), or cash-based public insurance or transfer programs (Danziger et al., 1982; Duncan et al., 1988; Moffitt, 1990; Hoffman and Duncan, 1995; Hoynes, 1997; Bobonis, 2011; Low et al., 2018; Persson, 2020).

While the research on changes to the regulation of private insurance show effects even for those of higher socioeconomic status, the studies of means-tested programs often hypothesize that only those of low socioeconomic status—proxied by education—are affected (Hampton and Lenhart, 2021; Chatterjee, 2024; Low et al., 2018). It is reasonable to assume that those most likely to have a low income (and qualify for such programs) at any given point in time also have a low education level. However, by conflating low income and low education in this way, we do not consider how the risk of serious negative income or job shocks—and public insurance against such shocks—affects the decisions of people who otherwise have a high expected income.

In this paper, we use the context of the Affordable Care Act's Medicaid expansion in the United States to explore the effects of public insurance on marital outcomes across the socioeconomic spectrum. This context is particularly useful for our analysis because of the close relationship between marital status and health coverage in the United States. Indeed, married individuals in the U.S. benefit from more options for health insurance coverage than their unmarried peers, reflecting a material benefit of marriage. For example, if an individual has employer-sponsored health insurance, this policy can provide coverage for a spouse if the spouse lacks or loses coverage. The expansion of Medicaid, a means-tested public health insurance program, provides an additional possible source of health insurance coverage for many unmarried adults. However, not everyone's spouse or pool of potential spouses are equally likely to have employer sponsored coverage: As a fringe benefit, it is positively correlated with the income and status of a position. As a result, the "marriage premium" for health insurance differs across the socioeconomic spectrum.

In practical terms, we use changes in the eligibility thresholds for Medicaid across states between 2011–2019 to capture variation in the likelihood of public health insurance eligibility. Using the American Community Survey (ACS), we construct a measure of the probability that unmarried individuals in a national sample and in a particular demographic group are eligible for Medicaid for each state and year, and we then look at the effects of that probability on marriage and divorce in the past year. This measure serves as a direct proxy for the probability of being eligible for Medicaid when unmarried. We regress transitions into and out of marriage on this measure to assess how changes in eligibility affect marital outcomes. The key identification assumption for this strategy is that policy changes in eligibility thresholds are uncorrelated with unobserved factors that may be related to marriage or divorce transitions.

Our empirical strategy is distinct from an IV-based strategy, where the proxy of coverage could be used as an instrument for the actual enrollment of individuals. Our research question concerns the effects of increases in the likelihood of becoming eligible for Medicaid if unmarried (even for individuals who are not currently eligible), and therefore an individual's actual enrollment is not our interest.

Empirically, we find that individuals with higher levels of educational attainment have decreased marriage rates in response to an increased likelihood of Medicaid eligibility, while there is no effect on marriage entry for those individuals with the lowest level of educational attainment. While the heterogenous results for marriage have not been identified in previous work, they are straightforward to understand given simple facts about health insurance in the U.S. First, there are extremely high potential costs to being uninsured in the U.S., meaning that even people with high expected income could reasonably be concerned with the risk of uninsurance. Accessing medical services while uninsured can cost thousands or tens of thousands of dollars and routinely puts households into debt and bankruptcy. Second, those individuals with the lowest level of educational attainment do not benefit from the aforementioned "marriage premium" for insurance, as this population has low levels of employer-sponsored health insurance. Since there is no health insurance benefit to marriage to begin with, expanding eligibility for unmarried individuals in this population would not affect their marriage decisions.

We address potential threats to our identification strategy, including migration to states with more generous eligibility thresholds and the differential selection of states into Medicaid expansion. To address selection into Medicaid expansion, we employ the "recentering" approach from Borusyak and Hull (2023) to condition for the expected likelihood of being covered by Medicaid outside of marriage, which plausibly leaves random variation in this likelihood and therefore leads to unbiased estimates of our treatment effect of interest. In robustness checks addressing these issues, the baseline results are unchanged.

Our results are fully reconcilable with recent related work by Hampton and Lenhart (2021) and Chatterjee (2024), who also study the effects of ACA Medicaid expansion on marriage and divorce but use different identification and sample restrictions. While they both find negative effects on entry into marriage for those with a high school degree or less, we also separately estimate effects for individuals with higher levels of education and given a more disaggregated definition of education, motivated by descriptive statistics on coverage rates that suggest that conditions for those who did not graduate high school are qualitatively different than they are for those who did.

In the second half of the paper, we present a search-based model of the marriage market to rationalize our empirical findings. In the model, individuals enjoy both pecuniary and non-pecuniary benefits of marriage. The latter might consider aspects such as love or the quality of a match, while the pecuniary benefits of marriage can be thought of as capturing the value of insurance or economies of scale. Individuals differ by type, which we think of as educational attainment. Individuals with higher educational attainment can bring a greater financial benefit to a partner through the potential for insurance. Like models where individuals learn about match quality over time, matched individuals can either decide to date or immediately enter marriage. It is more costly to separate when married than when unmarried.

This simple setup offers some striking implications about the tradeoffs between the pecuniary and non-pecuniary benefits of marriage. First, matches that deliver higher economic benefits will lead to marriage given lower levels of match quality than matches bringing lower economic benefits; the latter must have a high level of match quality for marriage to occur and be sustained. Second, because material benefits accordingly account for a greater share of the expected marital surplus for highly educated individuals, such individuals will respond more to changes in the economic value of their outside option—such as an expansion of public insurance—than those with less education. Specifically, they will become pickier about matches and delay marriage.

Building upon these insights, we simulate the general equilibrium effects of an expansion in public insurance. To do so, we calibrate an infinite horizon search model of the marriage market and use the parameters from this calibrated model to simulate the impact of an expansion in public insurance. First, we find a positive relationship between educational attainment and the baseline material benefits from marriage. Next, we find a monotone relationship between the baseline material benefits from marriage and the likelihood of delaying entry into marriage following an expansion of public insurance. This finding rationalizes our empirical results given the positive relationship between material surplus and educational attainment.

Our empirical and theoretical results show that the ACA Medicaid expansion did discourage marriage—not among those with the lowest socioeconomic status, but rather among those with higher socioeconomic status. These findings hint at a new hypothesis about why marriage rates are higher and divorce rates are lower for more compared to less educated individuals: More highly educated individuals might be content to accept a lower-quality match (compared to less educated individuals) because of the large economic benefits of marriage. Simultaneously, the larger economic benefits of marriage mean that these individuals are more sensitive to financial improvements in their outside options. Earlier results have demonstrated the value of shared assets in sustaining marriage for couples with positive net worth (Voena, 2015; Lafortune and Low, 2023), particularly in order to sustain investment in public goods. This paper suggests more generally that expected financial surplus in marriage, rather than shared assets specifically, also contributes to differential rates of entry into marriage, even in the absence of a public good such as children.

This potential implication of the model highlights the significance of our empirical and theoretical findings. Differences in marriage rates by educational status are one of the largest contributing factors to intra-household and intra-generational inequality in recent decades (Kearney, 2023). Having a clear understanding of what causes these differences is a necessary prerequisite to addressing the resulting growth in inequality. This paper takes several steps in this direction by demonstrating that the Affordable Care Act Medicaid expansion decreased entry into marriage for more educated individuals, while there was no effect or even some positive effects for the least educated individuals.

The paper is organized as follows. Section 2 provides the institutional background on the Affordable Care Act and the most relevant components of the Medicaid expansion. We lay out stylized facts about health insurance and marriage that motivate the key hypotheses in Section 3. A discussion of the dataset we use, our identification strategy and the empirical predictions are contained in Section 4. Section 5 describes the empirical results and the robustness checks pertaining to our empirical analysis. Section 6 presents the search-based model of the marriage market that we use as a theoretical framework to understand the direct and indirect implications of Medicaid expansion on the marriage market. Section 7 concludes the paper.

2 Institutional Background

In 2011, the U.S. Congress passed the Patient Protection and Affordable Care Act, a sweeping reform colloquially known as the ACA or Obamacare. In this section, we describe how options for health insurance changed as a result of the reform.

Prior to the ACA, there were effectively two options for health insurance for the under-65 population: employer-sponsored insurance (ESI) or Medicaid, a public insurance program for low-income and vulnerable groups.

Approximately 60% of the non-elderly population was covered by ESI before the reform. As a fringe benefit, ESI is positively correlated with earnings; lower-wage and part-time jobs often do not offer it. Important for our purposes, spouses could generally be covered by an employee's plan. Thus, if a married person loses his or her job and associated ESI, (s)he could have another option for coverage if his or her spouse also has ESI.

Medicaid is the federal name for what are largely state-administered programs; these have some federal requirements but substantial flexibility in setting coverage and eligibility thresholds. For example, in 2012, 17 states restricted eligibility for parents to levels under 50% of the federal poverty line, while 18 states set levels over 100% of the federal poverty line (Musumeci, 2012).

The ACA expanded Medicaid to reduce rates of uninsurance. The legislation mandated that states expand their Medicaid programs to cover all non-elderly adults under 138% of the federal poverty line (Foundation, 2010). The expansion was intended for national implementation in 2014, with an option for states to begin earlier. Individuals with ESI were not eligible for Medicaid before or after the reform.

However, upon passage of the ACA, 25 states sued the federal government in opposition to the Medicaid expansion. In 2012, the U.S. Supreme Court ruled that states could choose whether to participate in the Medicaid expansion (Musumeci, 2012). Six states expanded prior to 2014 (Foundation, 2012), and a further 19 states expanded in 2014. A number of additional states have expanded Medicaid in the years since, bringing the total to 33 states as of 2019 (not including the District of Columbia).

The uninsurance rate declined following implementation of the ACA, and in 2019 stood at approximately 10% for non-elderly adults. Most of this decline is attributable to Medicaid (Frean et al., 2017; Courtemanche et al., 2017).

While many studies of the effects of Medicaid focus on the simple classification between expansion and non-expansion states, this dichotomy masks considerable variation within and across these categories in the eligibility rules, in particular, in eligibility rules by parental status. We exploit this additional variation in our analysis (see Appendix C for details on the sources used to code eligibility rules).

An important feature of the Medicaid eligibility rules is that eligibility is determined by income threshold[1] and not assets (with a few exceptions, such as for elderly persons, who are not the focus in this paper). While public perception associates Medicaid with low socioeconomic status, people with high socioeconomic status experiencing transitory income shocks also make use of Medicaid, and not just those suffering long-term poverty. Over the past decade, individuals with a college education comprised approximately a tenth of Medicaid enrollees in any given year, and those with some college comprise approximately 15 to 20%, based on self-reported survey responses from the ACS. Unsurprisingly, these individuals are underrepresented compared to their population share, while those with lower levels of education are overrepresented, but these percentages nevertheless highlight that Medicaid is still relevant as a safety net program even for those with high educational attainment.

3 Stylized Facts on Health Insurance and Marital Status

In this section, we provide descriptive evidence about the relationship between marital status, health insurance, and educational attainment levels that is central to understanding the results of the paper.

Fact 1. A marital insurance gap exists.

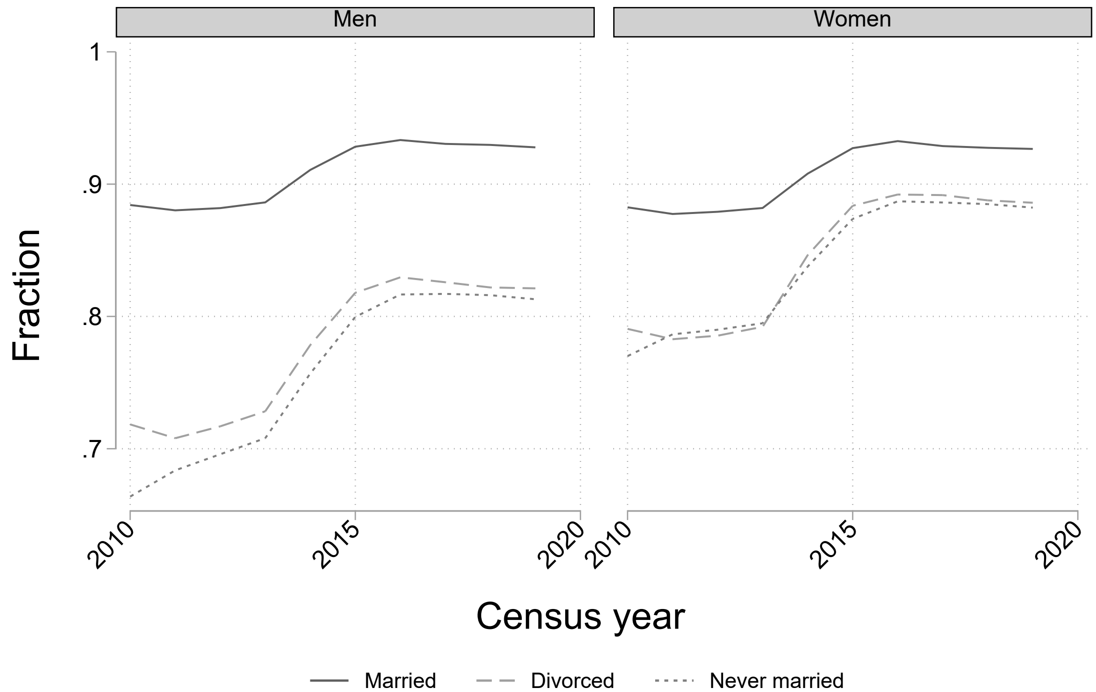
[1] We add that cohabitation is generally treated differently than legal marriage for the income test.

Our first stylized fact is that there is a substantial gap in average health insurance coverage rates between married and unmarried individuals. One explanation of this gap is the potential for spousal coverage via ESI. Figure 1 highlights the importance of this gap for the population as a whole, illustrating health insurance coverage rates for unmarried compared to married adults before and after the implementation of the ACA Medicaid expansion. In 2011, the gap in health insurance coverage between married and unmarried adult men was approximately 20 percentage points and that for women was approximately 10 percentage points.

Fact 2. Medicaid expansion decreased these marital insurance gaps.

Another finding is that these marriage insurance gaps shrank by nearly half around the year 2014, when a large number of states implemented the Medicaid expansion. Analysis of the ACA's effect on rates of uninsurance have shown that the Medicaid expansion was responsible for the majority of the increase in coverage of previously uninsured individuals (Frean et al., 2017; Courtemanche et al., 2017). These patterns suggest that Medicaid expansion may have meaningfully reduced the health insurance benefits of marriage for the population on average, as the health insurance options for unmarried individuals improved.

Figure 1: Health insurance coverage by marital status



This figure depicts the fraction of men and women that reported having any source of health insurance coverage, by marital status and year, using data from the American Community Survey waves 2011–2019 for adults between the ages of 18 and 64.

Fact 3. There is a "marriage penalty" for women with a low education level.

Though on average there is a "marriage premium" with respect to having health insurance, the coverage rates differ markedly by socioeconomic background as proxied by education. In the year prior to the passage of the ACA, the non-elderly adult uninsured rate was approximately 20%. Given Medicaid eligibility rules and the fact that ESI is generally only offered with full-time, well-paid positions, the uninsurance rate was markedly higher for individuals with less education: Prior to Medicaid expansion in 2011, approximately 40% of individuals with less than a high school education had no health insurance coverage, compared to 7% of those with a college education. Given that people often partner with others of a similar socioeconomic background, individuals also differ in likelihood that their (potential) spouse would have ESI that can be shared.

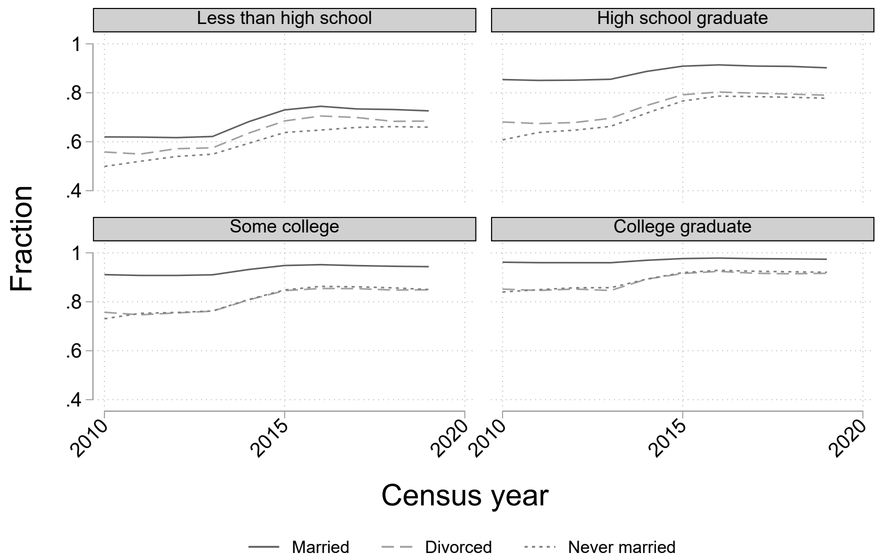
Figure 2 highlights how such differential access to ESI translates into differences in the "marriage benefit" of ESI for individuals with different levels of educational attainment. This figure shows the rates of health insurance coverage for married, divorced, and never married men and women across four levels of educational attainment between 2011 and 2019. For most groups, there is indeed a higher rate of coverage for married than for unmarried individuals indicative of a marriage benefit from ESI, as suggested by the population average depicted in Figure 1.

However, for women who did not graduate high school, their likelihood of having coverage while married is actually lower than their likelihood of being covered when single. As a result, there is a marriage penalty rather than a premium when it comes to health insurance for this group. This difference is likely attributable to low rates of ESI among their partners combined with the fact that many single women have custody of children and qualify for Medicaid under the more generous parental eligibility rules. Relatedly, among men who did not graduate high school, there are slightly higher rates of coverage for married men, but the gap is much smaller than that for all other men.

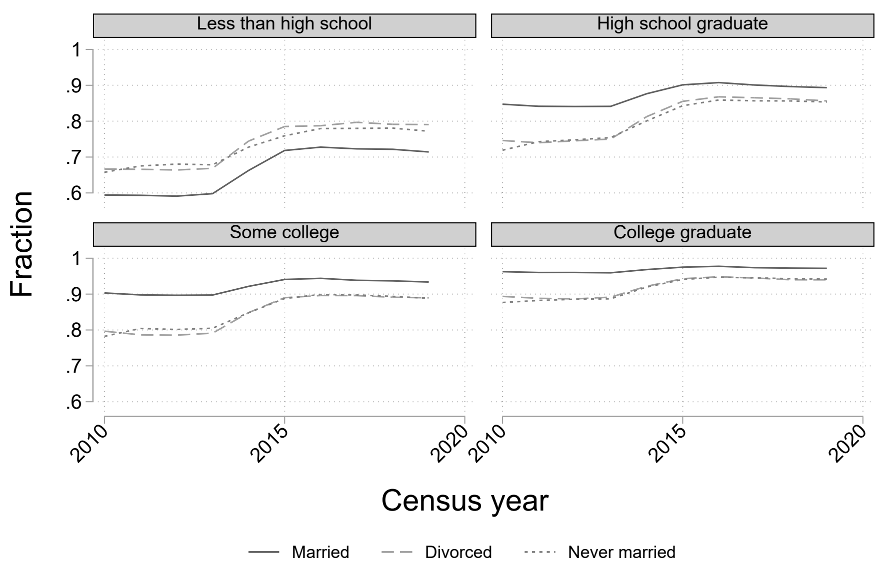
These differences across education highlight that prior to the Medicaid expansion, the structure of ESI offered an extra incentive to marry—but only for those whose partner or potential partners had a high likelihood of having ESI benefits through work. In turn, it suggests that changing the outside option for health insurance for unmarried individuals may differentially affect this subgroup.

Figure 2: Health insurance coverage by marital status and educational attainment

(a) Men



(b) Women



This figure depicts the fraction of men and women that reported having any source of health insurance coverage by marital status, year, and level of educational attainment, using data from the American Community Survey waves 2011–2019 for adults between the ages of 18 and 64.

4 Empirical Strategy

The goal of the empirical strategy is to estimate the average causal effects on marital outcomes associated with an increase in the probability of *becoming eligible* for Medicaid if unmarried. The intuition of studying this relationship is that an expansion of public health insurance may change the value of singlehood relative to that of marriage, which in turn may affect whether people decide to marry in the first place, what types individuals ultimately choose to marry, and some decisions of people who are already married. Concretely, we estimate the effect of contemporaneous increases in Medicaid eligibility on marriage in the last year (for previously unmarried individuals) and divorce in the last year (for already married individuals). We also estimate these effects by educational attainment level. In this section, we characterize the treatment of interest and discuss our key identifying assumptions. We then discuss the data and estimation.

4.1 Treatment of interest and identification

We are interested in how individuals' *likelihood of coverage* (not their actual coverage), assuming they were unmarried, affects their marital decisions and the marriage market equilibrium. Therefore, although actual coverage is binary, in this setting, the treatment of interest is continuous.

The likelihood that an individual will be eligible for coverage from Medicaid depends on two factors: their state's eligibility thresholds and the likelihood that their income will fall below this threshold. The eligibility threshold is fixed and can be observed at a given point in time. However, the likelihood that an individual's income will fall below a particular level is clearly neither given nor observed. We approximate it by treating the earnings distribution of unmarried individuals with similar characteristics as a proxy for an individual's potential earnings distribution.[2]

Concretely, we will use the observed earnings distribution for single individuals within specific demographic groups from a national sample and calculate the fraction within each group that falls below a given eligibility threshold. The earnings distribution is, by construction, orthogonal to local conditions and individual unobserved characteristics. The crucial identifying assumption is that state policy changes to eligibility thresholds are exogenous.

States might change the eligibility threshold in response to local economic or political conditions, which in turn may be correlated with marital outcomes. To address this concern, we only exploit variation in eligibility that stems from the choice to participate in the ACA Medicaid expansion, since this was a federal rather than state-level change. In practice, this means that we code all states with their 2011 eligibility levels and only update them if they participated in Medicaid expansion, ignoring other changes made at the state level alone.

The choice to use only variation from ACA expansions avoids the issue of capturing changes in eligibility that may be correlated with idiosyncratic state-level processes, but there is nevertheless a potentially important difference across the group of states that chose to expand versus not expand, especially in the early years after 2014: They were much more likely to be Democratic-leaning states, whether judging by the share that voted for Obama in his first election or by party control of the governorship or legislature. If Democratic-leaning states also have differential marriage rates compared to Republican-leaning states, then this would bias our estimates.

To address this additional threat to identification, we also employ the recenter-

[2]This approach is inspired by work using "simulated" instruments pioneered by Currie and Gruber (1996) and used by many others subsequently in studies of the effects of Medicaid. However, in those settings, the measure is generally used as an instrument for eligibility or actual coverage, which are both binary, whereas in our setting, it will serve as a proxy for the true continuous treatment.

ing approach from Borusyak and Hull (2023). They show that even when exposure to a treatment is non-random, unbiased estimates of the effects of the treatment can be obtained by conditioning on the expected likelihood of the treatment. Since the concern here is about the differential likelihood of expansion for Democratic-versus Republican-leaning states, we employ two methods for controlling the expected likelihood of a state choosing expansion: First, we partition states into strong Democratic/Republican or leaning Democratic/Republican based on the fraction in the state that voted for Obama in 2008. Next, we include fixed effects for these political categories interacted with individual year indicators. This strategy identifies the key effect of interest based on variation in Medicaid eligibility within these electoral categories. If one were still concerned that the variation among more and less Democratic states within these groups continues to bias our results, we also include a continuous measure of the fraction that voted for Obama interacted with the full set of year indicators.[3]

Beyond the relationship between political leanings and marriage rates, there remains one additional concern: Individuals may selectively move to states with higher eligibility, which could generate a relationship between marital outcomes and eligibility threshold if movers have systematically different marital outcomes. We also address this potential issue using two different strategies described as part of the estimation in Sec. 4.3.

[3]In practice, differential trends in marital rates for Democratic and Republican states seem to be less of a concern than one might imagine. While there are clear and persistent differences in marriage rate levels between Democratic-leaning states and Republican-leaning states—easily addressed by state fixed effects—there is no equally evident difference in trends over the time period of interest.

4.2 Data and variable construction

In this section, we describe our sources of data and the definitions of key variables. We use data from the ACS between 2011 and 2019, accessed via IPUMS (Ruggles et al., 2018). We include adults between the ages of 18 and 64, excluding those who would otherwise be covered by public insurance for the elderly.[4] The ACS includes information on current marital status, as well as information on changes in marital status from the previous year for select marital statuses, including the transitions into marriage and divorce. We also make use of demographic information and a constructed variable measuring income as a percent of the federal poverty level according to family income, family size, and the state in which the individual resides.[5]

To determine eligibility rules by state and year, we use information gathered from yearly reports between 2011 and 2019 by the Kaiser Commission on Medicaid and the Uninsured (see Appendix C for details). Eligibility levels for Medicaid vary within states by parental status, income as a percent of the federal poverty line, and work status. However, we use only the eligibility threshold levels for working individuals (which are higher than those for non-working individuals), as work status is endogenous, and the ACA expansion is intended to apply to a given level of income regardless of work status.

To calculate the key variable of interest, the probability that an individual would become eligible for Medicaid in the case in which (s)he is not married, we start with the national sample of unmarried individuals between 18 and 64 from the ACS for

[4]Those below the age of 26 also have the option of coverage via a parent's private insurance policy as a result of the ACA, which was implemented in 2009, but since this policy is uniform across the country over the full time period of the sample, its impact should be captured by age-group fixed effects

[5]There are distinct federal poverty lines for Alaska and Hawaii due to higher cost of living outside of the contiguous United States; for the 48 remaining states and the District of Columbia, there is only one standard, and it depends on family income and family size.

each year between 2011 to 2019 and randomly draw 10% of the ACS respondents along with the following characteristics. We partition the sample into groups by sex, 5-year age groups, and four levels of educational attainment (less than high school, high school graduate, some college, and college graduate). These characteristics are predictive of an individual's earnings potential and can be taken as predetermined in a given year for considering Medicaid eligibility. These partitioned groups give a proxy of an individual's expected earnings distribution that is not driven by local variation in earnings.

Table 1: Summary statistics for full sample

	Husbands		Wives	
	Mean	St. Dev.	Mean	St. Dev.
Age	41.58	(13.80)	42.20	(13.76)
More than high school	0.51	(0.50)	0.58	(0.49)
Parent	0.36	(0.48)	0.45	(0.50)
Medicaid Coverage	0.07	(0.09)	0.09	(0.10)
PrMedicaid	0.12	(0.13)	0.17	(0.18)
Married	0.49	(0.50)	0.52	(0.50)
Divorced	0.10	(0.30)	0.12	(0.33)
Marriage Rate (Unmarried in t-1)	0.04	(0.20)	0.04	(0.20)
Divorce Rate (Married in t-1)	0.02	(0.13)	0.02	(0.13)
Observations	8315501		8505164	

This table provides summary statistics for data from the American Community Survey. The sample includes all adults ages 18–64 from 2011–2019.

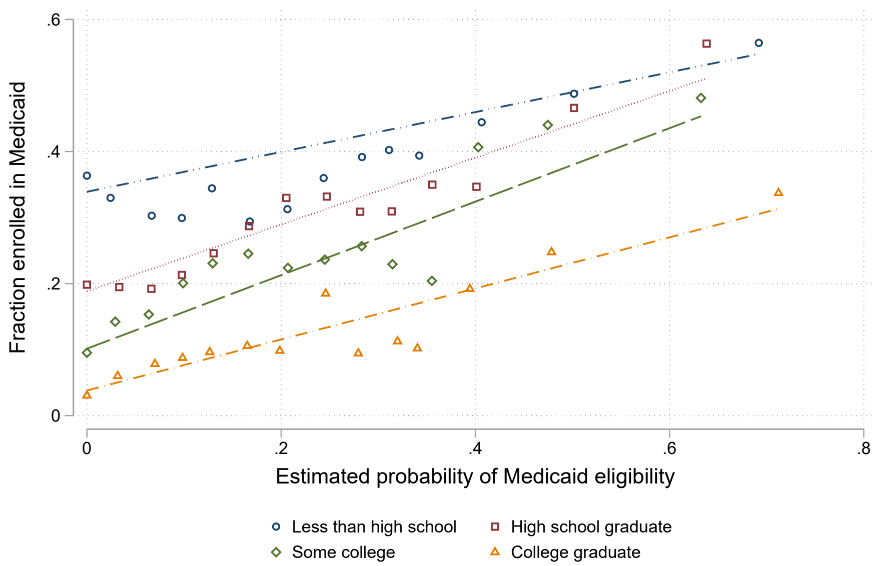
For each demographic group in a given year, we calculate the fraction that would be eligible under each state's eligibility rules based on their income as a percent of the federal poverty line and their parental status, with the contribution of each individual being weighted by person-specific survey weights. This fraction, $PrMedicaid_{d,s,t}$ (where d indicates the demographic group), is our proxy for the unobserved probability of Medicaid coverage when not married.

Table 1 provides descriptive statistics for the full sample of adults. Men and women have a mean age of 41.6 and 42.2, respectively. Women are nearly 10 percentage points more likely to be classified as parents (defined as having a child under 18 in the household), which reflects the fact that in the case of parental separation, mothers most often have primary custody. Likewise, women are slightly more likely to be enrolled in Medicaid (0.09 compared to 0.07) or eligible for Medicaid when not married (0.17 to 0.12). The annual marriage rate is approximately 4 per 100 married couples, while the annual divorce rate is approximately 2 per 100 couples.

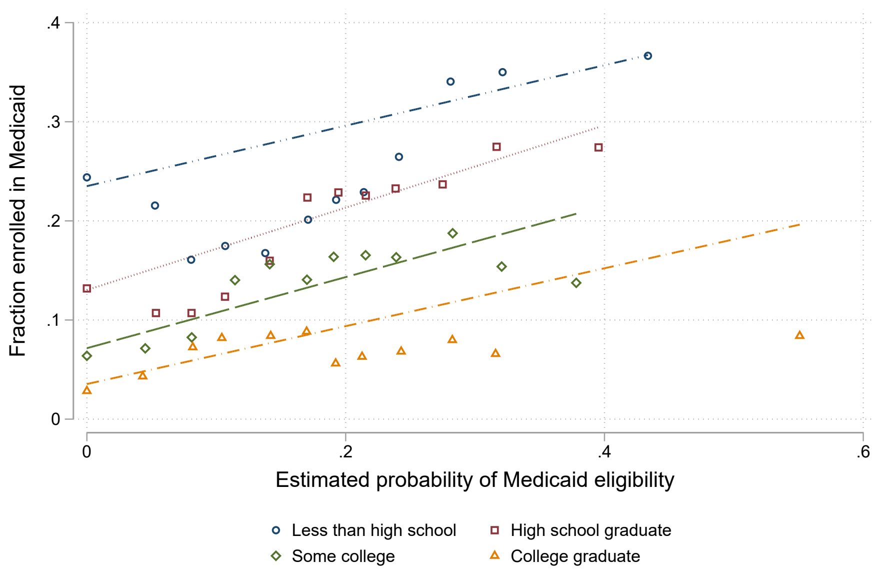
Although we are not employing a two-stage estimation strategy, since we are not interested in the direct effect of enrollment, we can nevertheless provide support for the relevance of our key variable $PrMedicaid_{d,s,t}$ by showing that higher values are also correlated with higher levels of Medicaid enrollment for unmarried individuals across all levels of educational attainment. This correlation is illustrated in the binned scatterplots in Figure 3.

Figure 3: Fraction of unmarried adults enrolled in Medicaid by PrMedicaid

(a) Women



(b) Men

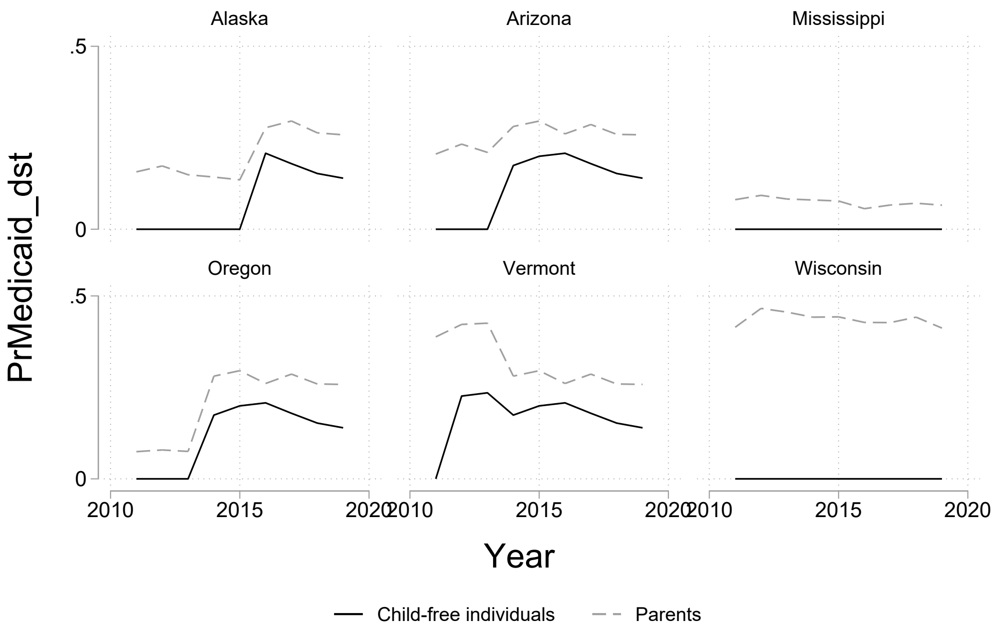


The figure shows a binned scatterplot for individuals who are not married from the American Community Survey, years 2011–2019. The horizontal axis depicts the average value of $\text{PrMedicaid}_{d,s,t}$ for each 5% bin in this sample of respondents. The vertical axis correspondingly shows the average level of reported Medicaid enrollment for each bin.

To provide some intuition for the identifying variation exploited here, Figure 4 graphs $PrMedicaid_{d,s,t}$ for two demographic groups, showing the estimated eligibility for six states for both parents and childless individuals who are female, are between the ages of 30 and 35, and have attended but not graduated from college. The states chosen highlight that we make use of additional variation beyond the classification of expansion and non-expansion states. In particular, states vary in the initial eligibility threshold for parents, with changes that differ even upon expansion. For example, Vermont lowered the eligibility threshold for parents (which had been above 138% of the federal poverty line), while Alaska, Oregon, and Arizona all increased eligibility for parents upon expansion, but with differing magnitudes. The timing of expansion decisions also differed, providing additional variation. Even among non-expansion states, parental eligibility rates vary, as evident from a comparison of Mississippi and Wisconsin, although as non-expansion states, they exhibit no variation over time[6]

[6]In practice, Wisconsin made a state-level change to its Medicaid program that made their program nearly as generous in terms of eligibility threshold as the ACA-mandated levels, but they did not accept federal funding to do so.

Figure 4: Eligibility among women 30–35 years old, some college



The figure depicts the $\text{PrMedicaid}_{d,s,t}$ for women aged 30–35 years who have attended but not graduated from college across selected states. $\text{PrMedicaid}_{d,s,t}$ is constructed by calculating the fraction of individuals with the corresponding characteristics in a fixed national sample that would be eligible for Medicaid under the state's rules in a given year. The data are taken from the American Community Survey, years 2011–2019.

4.3 Estimation

In this section, we explain our estimation given our identification assumptions. We estimate effects first on transitions into marriage and then on transitions into divorce.

4.3.1 Empirical testing for effects on new marriages

Intuitively, expanded insurance affects individuals' outside options relative to marriage. In the first part of our analysis, we consider whether the Medicaid expansion affects marriage rates.

As unmarried individuals' outside options improve, we expect that they may become pickier about potential spouses and become less likely to marry.

Using a linear probability model estimated by OLS, the main equation that we will estimate associated with this prediction is

$$\text{MaritalTransition}_{i,s,t} = \eta \text{PrMedicaid}_{d,s,t} + \gamma_d + \delta_t + \mu_s + \mathbf{X}_{i,s,t}\beta + \varepsilon_{i,s,t} \qquad (1)$$

The equation is estimated for women and men separately. The subscript d indicates a demographic group, s is the state in which they reside, and t is the year. We consider two dependent variables: first, whether an individual married in the past year, $\text{MarrPastYr}_{i,s,t}$, with the sample in this case including only individuals who were unmarried the previous year. The second dependent variable is whether an individual divorced in the past year, $\text{DivPastyr}_{i,s,t}$, with this sample then including individuals who were among the married population in the prior year.

The key variable of interest for both dependent variables is $\text{PrMedicaid}_{d,s,t}$, which characterizes the probability that an individual would be eligible for Medicaid outside of marriage. Every regression includes a set of fixed effects for each demographic group indicated by γ_d (e.g., each group defined by age, parental status, and education), year fixed effects δ_t, and state fixed effects μ_s. Additionally, $\mathbf{X}_{i,s,t}$ represents a vector of supplemental control variables, which includes indicators for number of children (1, 2, or 3+), as well as further time-varying covariates of economic conditions for demographic groups, including the mean log wage, the mean unemployment rate, and the mean labor force inactivity rate by demographic group and state.

Next, we consider two ways of applying the recentered instruments strategy Borusyak and Hull (2023). First, we include four political categorizations of states: whether they were strongly or weakly Democratic-/Republican-leaning in the 2008

U.S. presidential election.[7] We interact these four political categories with a full set of year indicators to control for differential propensities to select into Medicaid expansion over time. To provide an even more flexible approach, we also implement an additional specification whereby we use the fraction of the state that voted for Obama interacted directly with year indicators.

Finally, we also consider that selective migration to states that increase Medicaid eligibility may be a potential source of bias. We test for this bias in two ways, first by adding a control for the fraction of individuals in a demographic group and state that moved to that state in the last year, and second, by excluding all recent movers from the sample. Standard errors are clustered at the state level in all specifications.

When considering effects specific to educational subgroups, we interact $PrMedicaid_{d,s,t}$ with the categorical variable of interest and calculate the total marginal effects for each subgroup. The specifications applied are the same as those used for the overall average effect.

5 Results

We now turn to a discussion of the empirical findings. When considering overall population averages, we find a negative effect on marriage and no effect on divorce. When distinguishing effects by individuals' differing levels of educational attainment, we find significant and negative effects on marriage rates for those with higher levels of educational attainment and no effect on marriage for those with the lowest level of educational attainment.

We start with the results for entry into marriage. Table 2 presents results for

[7] To generate cutoffs, we partition first at the 50% mark and then split each of the remaining number of states in half, which yields cutoffs off greater than 57% voting for Obama for strongly Democratic and less than 43% voting for Obama for strongly Republican.

the population as a whole in the rows labeled "Overall Effect", while the rows for each level of education present the total marginal effect of increased probability of Medicaid eligibility for that subgroup[8].

The overall average effect on marriage of an increase in the probability of Medicaid eligibility is negative for both men and women, with point estimates at approximately -0.008 for men and -0.004 for women, differing only slightly across specifications. It is not statistically significant for women overall.

When disaggregating this average effect by educational attainment level, we observe clear differences. For those with the lowest level of education, the key coefficient is not significant and close to zero, with point estimates ranging from -0.001 to 0.002 for men and 0.000 to 0.003 for women. For all other educational levels and across specifications, the coefficients are larger in magnitude than in the overall average estimate, ranging from -0.0006 (for women with a high school degree) to -0.042 (for women with a college degree).

The confidence intervals for those with a high school degree and some college are largely overlapping, while the confidence intervals for those with a college degree are partially overlapping with the other two, although they have much more negative point estimates. These estimates suggest that when faced with increases of similar magnitudes in the likelihood of Medicaid eligibility, those with higher levels of education are more likely to respond by avoiding or delaying marriage than those with lower levels of education.

Implementing the recentering approaches to address differential partisan selection into Medicaid expansion and applying strategies to adjust for a confounding role of migration, either by controlling for the share of recent movers or by excluding recent

[8]To be clear, the coefficients for the overall average effect are from one regression, while the coefficients for the total marginal effects by education are estimated simultaneously from one additional regression per specification

movers altogether, do not markedly change the results, suggesting that these are not, in fact, substantial sources of bias.

Table 2: The effect of the probability of Medicaid eligibility on marriage rates

	Baseline		Recentering		Migration	
	(1)	(2)	(3)	(4)	(5)	(6)
	Men					
Overall Effect						
PrMedicaid_{dst}	-0.008	-0.008	-0.010	-0.009	-0.010	-0.010
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Effect by Edu.						
Edu $<$ HS	0.002	0.002	0.000	0.000	-0.001	-0.002
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu = HS	-0.015	-0.015	-0.019	-0.018	-0.019	-0.019
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Edu = Some College	-0.012	-0.010	-0.014	-0.013	-0.014	-0.013
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Edu $\geq$ College	-0.028	-0.028	-0.035	-0.033	-0.034	-0.036
	(0.012)	(0.012)	(0.013)	(0.012)	(0.013)	(0.013)
Observations	3565222	3565222	3565222	3565222	3565222	3431430
	Women					
Overall Effect						
PrMedicaid_{dst}	-0.004	-0.004	-0.004	-0.004	-0.004	-0.005
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Effect by Edu.						
Edu $<$ HS	0.003	0.003	0.002	0.002	0.002	0.000
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu = HS	-0.007	-0.006	-0.008	-0.007	-0.008	-0.008
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu = Some College	-0.010	-0.010	-0.012	-0.011	-0.011	-0.011
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu $\geq$ College	-0.038	-0.039	-0.042	-0.041	-0.042	-0.042
	(0.012)	(0.012)	(0.011)	(0.011)	(0.011)	(0.011)
Observations	3721213	3721213	3721213	3721213	3721213	3586134
Economic Controls		✓	✓	✓	✓	✓
Recentering			Discrete	Cont.	Discrete	Discrete
Migration					Control	Exclude

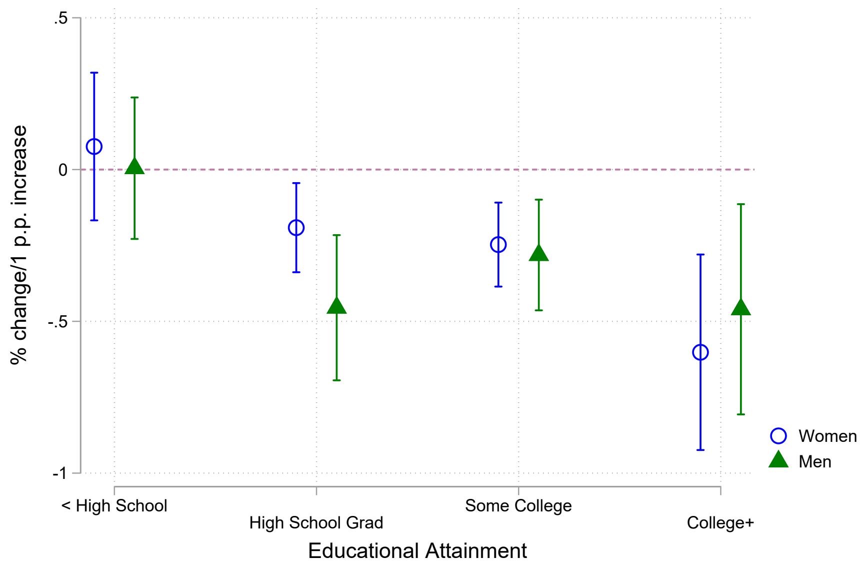
Results are estimated by OLS using adults ages 18 to 64 from the American Community Survey, waves 2011–2019. Whether an individual was married in the prior year is regressed on the probability of Medicaid eligibility given the individual's demographic characteristics. Each regression controls for the fully interacted set of demographic group fixed effects (by 5-year age group, four levels of educational attainment, and parental status), as well as for year and state fixed effects. Economic controls include time-varying covariates for the demographic group's unemployment rate, mean log wage, and labor force inactivity rate; discrete recentering indicates the inclusion of four political categories (by level of support for Obama in the 2008 election) interacted with a full set of year indicators, while continuous recentering indicates that the share of the vote for Obama is directly interacted with year indicators; the strategy to control for migration adds a control for the share of individuals who moved into the relevant state in the past year per demographic group, while "exclude" indicates that these same individuals were dropped from the regression.

There is one important caveat when comparing the effect sizes across the educational attainment groups. First, the baseline levels of entry into marriage differ across educational levels, with those with higher educational attainment also having higher levels of entry into marriage. To provide another way of comparing these estimates, we create a figure with scaled effect sizes by educational attainment level (using the results from Col. 3 in Table 2). Specifically, we rescale the estimated coefficient for each subgroup by dividing each coefficient by the group's baseline marriage rate, effectively yielding a semi-elasticity that expresses the percent change in the marriage rate in response to a 1 percentage point increase in the probability of Medicaid eligibility.

This comparison yields more moderate differences. Whereas the coefficients for college-educated individuals were three to five times larger than the coefficients for those with less education, the scaled effect size when accounting for the differences in baseline marriage rates differ only by approximately a factor of two. Nevertheless, those with a college education still have clearly larger responses. The scaled effect sizes suggest that for every 1 percentage point increase in the likelihood of Medicaid eligibility, those with a high school degree or some college education experience a reduction of about a quarter to a half a percent reduction in entry into marriage, whereas those with a college education see an approximately a half a percent reduction in marriage entry.

Considering why those with less than a high school education have an anomalous result, we refer back to the fact that women in this group are the only cohort who had higher levels of insurance coverage while unmarried than married prior to the passage of the ACA. For this group, there was no marriage premium in the first place, while for all other groups, Medicaid expansion reduced a (positive) marriage premium associated with health insurance. Together, these results on marriage suggest that individuals in marginal relationships postpone or avoid marriage, and that for this particular Medicaid expansion, this effect is concentrated among those with higher rather than lower levels of educational attainment, particularly for women.

Figure 5: The effect on marriage rates of an increased probability of Medicaid eligibility by educational attainment



Including individuals between the ages of 18 and 64 from the 2011 to 2019 waves of the ACS, this graph presents scaled coefficients from a regression of marriage in the last year on the probability of Medicaid eligibility for individuals in a demographic group defined by five-year age groups, parental status, and educational attainment, where the probability of Medicaid eligibility is interacted with the educational attainment group to yield education-group specific total marginal effects. Each regression includes fixed effects for demographic group, year, state, and the demographic group's time-varying economic conditions as well as the state governor's party affiliation. Coefficients are divided by each group's baseline marriage rate to yield the percent change in marriage rate for a 1 percentage point increase in $\text{PrMedicaid}_{d,s,t}$.

We have also posited that eligibility at the time of marriage may affect divorce rates. Table 3 presents results pertaining to this hypothesis, with an analogous structure to that used for the results for marriage. However, in all specifications and for all educational subgroups, the coefficient on $\text{PrMedicaid}_{d,s,t}$ is not significantly different from zero.

Table 3: The effect of an increased probability of Medicaid eligibility on divorce rates

	Baseline		Recentering		Migration	
	(1)	(2)	(3)	(4)	(5)	(6)
			Men			
Overall Effect						
PrMedicaid_{dst}	-0.003	-0.002	-0.002	-0.002	-0.002	-0.003
	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Effect by Edu.						
Edu $<$ HS	0.000	0.001	0.001	0.001	0.001	0.001
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu = HS	-0.007	-0.006	-0.006	-0.006	-0.006	-0.007
	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Edu = Some College	-0.006	-0.006	-0.006	-0.006	-0.005	-0.006
	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)	(0.008)
Edu $\geq$ College	-0.015	-0.016	-0.017	-0.017	-0.016	-0.016
	(0.010)	(0.010)	(0.011)	(0.011)	(0.011)	(0.011)
Observations	3980472	3980472	3980472	3980472	3980472	3882862
			Women			
Overall Effect						
PrMedicaid_{dst}	-0.002	-0.002	-0.002	-0.002	-0.002	-0.002
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Effect by Edu.						
Edu $<$ HS	0.000	-0.001	-0.001	-0.001	-0.001	-0.001
	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)
Edu = HS	-0.002	-0.002	-0.002	-0.002	-0.002	-0.002
	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu = Some College	-0.005	-0.005	-0.005	-0.005	-0.005	-0.005
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
Edu $\geq$ College	0.000	-0.001	-0.002	-0.002	-0.002	-0.004
	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)	(0.007)
Observations	4421027	4421027	4421027	4421027	4421027	4313042
Economic Controls		✓	✓	✓	✓	✓
Recentering			Discrete	Cont.	Discrete	Discrete
Migration					Control	Exclude

Results are estimated by OLS using adults ages 18 to 64 from the American Community Survey, waves 2011–2019. Whether an individual was divorced in the prior year is regressed on the probability of Medicaid eligibility given the individual's demographic characteristics. Each regression controls for the fully interacted set of demographic group fixed effects (by 5-year age group, four levels of educational attainment, and parental status), as well as year and state fixed effects. Additional controls include time-varying covariates for the demographic group's unemployment rate, mean log wage, and labor force inactivity rate and the state governor's party affiliation.

The lack of an economically meaningful positive effect on contemporaneous divorce perhaps goes against common intuition. Hampton and Lenhart (2021) find an increase in divorce stocks, but this change could result from a decrease in marriage rates, since it may be that divorcees remain divorced; they also do not find a direct effect on divorce flows. Nevertheless, it is possible that we are simply unable to identify the true effects of Medicaid expansion on divorce, since divorce can take years from the decision to a legal change in status, whereas (avoiding) marriage is an immediate decision.

6 Theoretical Framework

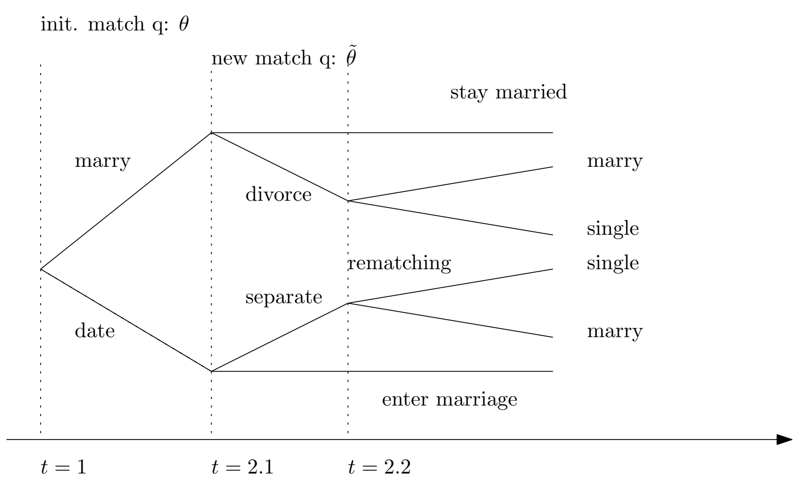
The empirical analysis highlights the broad effects of increases in Medicaid eligibility (public insurance) on the marriage market. The aim of this section is to present a stylized model of the marriage market that can rationalize these findings. In terms of the findings, we found an overall decrease in the marriage rate and, in particular, a notably stronger proportional response for more highly educated individuals. The key contribution of our model is that it provides an explanation for this latter result. The main channel is the tradeoff between material and immaterial benefits in marriage. In particular, with a higher initial material value of marriage, more highly educated couples are *less* picky about quality than less educated individuals. In turn, because the material value of marriage for less educated individuals may be low or even negative, a decrease in the relative value of marriage through public insurance expansion is predicted to have little impact on their marital decisions, as those are mainly driven by immaterial benefits from marriage.

6.1 Setting

We consider a simple two period model of the marriage market. We consider a couple's formation between opposite-sex individuals. We assume there is a continuum of men (m) and women (f), each with measure one. Individuals discount the future by a factor $\delta \in (0,1)$. Furthermore, individuals are divided into types, which will be indicated by indices $e \in E$. For the purposes of the current paper, and to remain in line with the empirical analysis, we interpret e as the educational attainment of a particular individual. In this case, $E = \{H, L\}$, where H refers to a 'high' level of educational attainment, which we consider as having attended college, and L represents having obtained less than some college education. The distribution over types is given by $\left(\gamma^m, \gamma^f\right)$, with $\gamma^g = \left(\gamma_H^g, \gamma_L^g\right), g = m, f$.

Timing in the model is represented in Figure 6. In period 1, each individual is randomly matched with an opposite-sex (potential) partner. After they have been matched, the initial match quality θ is revealed, drawn from a distribution F. After the revelation of match quality, they can then decide whether to date their partner or immediately enter marriage.[9]

Figure 6: Timeline model



In period 2, those who married in period 1 can either divorce or remain married. Those who were dating can either enter marriage with their partner or separate.

[9]We opted for dating as the closest substitute choice to marriage in $t = 1$, given that we believe this approximates the real decision of individuals upon an expansion of their outside value in the context of the ACA, i.e., whether or not to enter legal marriage.

Furthermore, those individuals who decided to separate or divorce from their period 1 partners can match with a new partner from the pool of singles. Upon meeting a single, they can then decide whether to remain single or to marry the individual they meet in that round.

The utility flows in marriage between a man with educational attainment e and a woman who has educational attainment e' are given by $M_{e,e'}$. We normalize the utilities of individuals outside of marriage (either while dating or being single) to zero. We are assuming perfectly transferable utility, which is relatively standard in much of the related literature (e.g., Goussé et al. (2017)).

Match qualities are not fixed, reflecting that the quality of relationships can change over time. At the start of period 2, a new match quality, $\tilde{\theta}$, is drawn from a conditional distribution $G\left(.|\theta\right)$. Since it is conditional on the first-period match quality, we allow for correlation in match quality levels between the two periods. If a couple starts with a high level of match quality, it is more likely to also be high in the next period. Individuals who choose to divorce at the start of period 2 are rematched randomly with another single individual, and after meeting, they draw a match quality from the same distribution F as used in period 1.

A key difference between dating and marriage is in the cost of separation: While separating from one's partner is costless after dating, we assume that individuals who divorce incur a cost equal to $\kappa > 0$. This cost of divorce means that married couples will stay together in the face of negative shocks even at a quality level that would cause dating couples to separate. In turn, this greater resilience to negative quality shocks means that married couples also have greater expected benefits from intra-household insurance than dating couples.

6.2 Equilibrium analysis

We now solve the model, working by backwards induction. Two events can occur in period two, yielding two subperiods: In $t = 2.1$, individuals choose whether to remain with a partner from period 1. In $t = 2.2$, individuals who choose to separate or divorce are rematched to another single. Working backwards, we start with $t = 2.2$, where individuals have been rematched after separation or divorce. A couple consisting of a type e man matched with a type e' woman will choose to marry if and only if

$$M_{e,e'} + \theta \geq 0,$$

which occurs with probability $\hat{\pi}_{e,e'} = 1 - F\left(-M_{e,e'}\right)$. In $t = 2.2$ the probability of meeting a partner with a specific level of educational attainment is determined by the composition of the pool of singles resulting from the separation and divorce decisions at $t = 2.1$. This composition is characterized by $s^m = (s^m_H, s^m_L)$ and $s^f = \left(s^f_H, s^f_L\right)$, which denote the measure of single men and women by educational attainment. Similarly, $S^m = s^m_H + s^m_L$ and $S^f = s^f_H + s^f_L$ are the total number of single men and women, respectively. The probability that a single man with educational attainment e meets a single woman with an education level e' is then given by

$$\alpha_{e,e'} = \frac{s^m_e}{S^m}\frac{s^f_{e'}}{S^f},\ e, e' \in E.$$

Using these meeting probabilities and assuming that marital surplus is split equally between prospective spouses, we can compute the continuation values for single men and women across educational types. We denote these by $\mathcal{W}_{e,0}$ for a single man with education level e and by $\mathcal{W}_{0,e'}$ for a single woman with educational attainment e'.[10]

Based on these continuation values, in period $t = 2.1$, individuals decide whether to stay with their partner from period $t = 1$. We must distinguish two cases, depending on whether the couple was dating or married in period $t = 1$. First, consider a marriage between a man of type e and a woman of type e'. These individuals will remain married when

$$M_{e,e'} + \tilde{\theta} \geq \mathcal{W}_{e,0} + \mathcal{W}_{0,e'} - \kappa, \tag{2}$$

which states that the utility flow of remaining married (on the left hand side) is at least as large as the overall utility both spouses could obtain if they divorced, netting out the cost of divorce. Assuming that this couple's initial match quality was equal to θ, the probability of remaining married is then given by

$$\tilde{\pi}^{mar}_{e,e'}(\theta) = 1 - G\left(\underline{\tilde{\theta}}^{mar}_{e,e'}|\theta\right), \tag{3}$$

where we have defined $\underline{\tilde{\theta}}^{mar}_{e,e'} = -M_{e,e'} + \mathcal{W}_{e,0} + \mathcal{W}_{0,e'} - \kappa$, the lowest level of match quality at which individuals of the given educational type prefer to remain married. A similar analysis can be done for those individuals who were dating in period $t = 1$: A man of type e and a woman of type e' who were dating will choose to marry in $t = 2$ if

$$M_{e,e'} + \tilde{\theta} \geq \mathcal{W}_{e,0} + \mathcal{W}_{0,e'}. \tag{4}$$

The probability of a transition from dating to marriage is therefore given by $\tilde{\pi}^{dat}_{e,e'}(\theta) = 1 - G\left(\underline{\tilde{\theta}}^{dat}_{e,e'}|\theta\right)$, where $\underline{\tilde{\theta}}^{dat}_{e,e'} = -M_{e,e'} + \mathcal{W}_{e,0} + \mathcal{W}_{0,e'}$.

Clearly, equation (2) is very similar to (4), with the difference being in the ad-

[10] We refer to Appendix D.1 for the expression of these continuation values.

ditional term representing the cost of divorce. Conditional on the types of partners and the initial match quality, it can easily be seen that the probability of remaining together in period 2 is larger when the couple married in $t = 1$, that is, $\tilde{\pi}^{dat}_{e,e'}(\theta) \leq \tilde{\pi}^{mar}_{e,e'}(\theta)$. Intuitively, due to the presence of the additional barrier to separation when married in $t = 1$, individuals can afford larger adverse shocks to match quality and still prefer to stay together rather than divorce.

Those couples who enjoy larger material benefits from being married—larger $M_{e,e'}$—also have an increased likelihood of remaining together in addition to greater incentives to enter marriage.

Moving to the decision in period $t = 1$, individuals facing a randomly drawn potential partner have to decide whether to immediately enter marriage or first date their prospective partner. Consider again a man of type e and a woman of type e' who are matched with each other in $t = 1$. The expected surplus of such a couple if they marry is given by

$$\mathcal{S}^{mar}_{e,e'}(\theta) = M_{e,e'} + \theta + \delta\mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}^{mar}_{e,e'}, 0\right\}\right], \tag{5}$$

Similarly, the expected surplus of dating is equal to

$$\mathcal{S}^{dat}_{e,e'}(\theta) = \delta\mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}^{dat}_{e,e'}, 0\right\}\right], \tag{6}$$

where $\mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}^{mar}_{e,e'}, 0\right\}\right]$ ($\mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}^{dat}_{e,e'}, 0\right\}\right]$) denotes the expected continuation value for the second period for married and dating couples, respectively, which is determined by the greater of either the surplus of staying together with the first period partner or that from separating and finding a new match.

The couple will then decide to marry immediately in case

$$\mathcal{S}_{e,e'}^{mar} - \mathcal{S}_{e,e'}^{dat} \geq 0. \tag{7}$$

This condition can be equally characterized by an acceptance rule of the following form:

$$\theta \geq \underline{\theta}_{e,e'},$$

that is, only couples with a sufficiently large match quality will immediately enter marriage; otherwise, they will first choose to date each other and then make a final decision on whether to enter marriage or try to find a new prospective partner in period $t = 2$. The probability of this occurring is given by $\pi_{e,e'} = 1 - F\left(\underline{\theta}_{e,e'}\right)$.

In practical terms, solving the model requires finding a fixed point. Given a distribution of singles over gender and educational attainment, $s = \left(s^m, s^f\right)$, one can compute the expected continuation values for becoming single in $t = 2$, which in turn determine the reservation match qualities, $\underline{\tilde{\theta}}_{e,e'}^{mar}$ and $\underline{\tilde{\theta}}_{e,e'}^{dat}$ for all $e, e' \in E$, as well as the period 1 critical match qualities determining the marriage decisions, $\underline{\theta}$. Conversely, the number of singles is given by those reservation match qualities. For example, consider (8). The total number of single men with a high level of education can be computed as those high education-type men who are either dating or married to any woman (irrespective of their education level) and subsequently separated in $t = 2.1$. A similar calculation holds true for women and less educated men. In sum,

$$s^m_e = \gamma^m_e \sum_{e'} \gamma^f_{e'} \left(\pi_{e,e'} \mathbb{E}\left[1 - \tilde{\pi}^{mar}_{e,e'}(\theta) \,|\theta \geq \underline{\theta}^*_{e,e'}\right] + (1 - \pi_{e,e'}) \mathbb{E}\left[1 - \tilde{\pi}^{dat}_{e,e'}(\theta) \,|\theta < \underline{\theta}^*_{e,e'}\right]\right) \tag{8}$$

$$s^f_{e'} = \gamma^f_{e'} \sum_{e} \gamma^m_{e} \left(\pi_{e,e'} \mathbb{E}\left[1 - \tilde{\pi}^{mar}_{e,e'}(\theta) \,|\theta \geq \underline{\theta}^*_{e,e'}\right] + (1 - \pi_{e,e'}) \mathbb{E}\left[1 - \tilde{\pi}^{dat}_{e,e'}(\theta) \,|\theta < \underline{\theta}^*_{e,e'}\right]\right) \tag{9}$$

We can now formalize a marriage market equilibrium for this framework:

Definition 6.1 (marriage market equilibrium)**.** *The marriage market equilibrium consists of a distribution of singles by gender and educational attainment* $s^* = \left(s^{m*}, s^{f*}\right)$ *and reservation match qualities* $\underline{\theta}^*_{e,e'}, \underline{\tilde{\theta}}^{mar,*}_{e,e'}, \underline{\tilde{\theta}}^{dat,*}_{e,e'}$ *such that*

- *Given* s^**, the reservation match qualities for both periods,* $\underline{\tilde{\theta}}^{mar,*}_{e,e'}, \underline{\tilde{\theta}}^{dat,*}_{e,e'}, \underline{\theta}^*$ *resp. solve* (2)*,* (4) *and* (7) *with equality.*
- *The number of singles by gender and educational attainment are solutions to* (8) *and* (9)*, given the reservation match qualities.*

It can be shown that a marriage market equilibrium always exists in our setting, which we state formally in the following:

Proposition 6.2 (Existence of marriage market equilibrium)**.** *If* $\delta \in [0, 1)$ *and* $\kappa \in [0, \epsilon)$ *for some small* $\epsilon > 0$*, a marriage market equilibrium exists.*

Proof. See Appendix. □

6.3 Effects of public insurance

We now consider the effects of the introduction of public insurance on the marriage market. In practice, we want to consider a case where the (expected) public transfers are larger for individuals when they are single compared to when they are (legally) married, which is consistent with the fact that most public insurance schemes are means tested at the household level. For example, a job loss for an individual almost certainly qualifies them for public insurance, while it may not do so for a married couple. To capture this in the context of our model, we will assume that public insurance corresponds to a reduction in the (net) material gains from marriage, i.e., $\Delta M_{e,e'} \leq 0$, with a strict inequality for at least one match type. We consider both the partial equilibrium effects as well as the general equilibrium effects, using a calibration exercise for the latter.

6.3.1 Partial equilibrium effects

First, we consider the implications of public insurance in partial equilibrium.

To gain some intuition, let us focus on one specific match type, e, e', and consider the introduction of public insurance, which is then a decrease in $M_{e,e'}$. Keeping the number of singles fixed, in period $t = 2.2$, those individuals who separated from their partner in $t = 2.1$ will now be less likely to marry their proposed match, given that the material benefits of marriage have decreased. At the same time, the average match quality of the marriages that are formed are higher.

Moving backwards in time, in period $t = 2.1$, when individuals must decide whether or not to stay together with their current partner (i.e., separate or marry), the decrease in material marital benefits and increase in (expected) average match qualities for new matches would imply an increase in the reservation match quality, thereby making it more likely that couples formed in period $t = 1$ separate. However, counteracting this is the fact that the probability of forming a new couple in period $t = 2.2$ decreases, which decreases the outside option for individuals in period $t = 2.1$ and thereby makes them less demanding with regards to the level of match quality required to stay with their partner.

Consequently, the (expected) effect on $\tilde{\underline{\theta}}^{mar}_{e,e'}$ and $\tilde{\underline{\theta}}^{dat}_{e,e'}$ is ambiguous. In turn, the effect on the divorce (or separation) probability is also ambiguous. This latter aspect is in line with our empirical findings, given that we find little evidence of a (net) effect on divorce for already married couples.

We now turn to the initial period $t = 1$, when individuals enter the dating and marriage market. Similar to the analysis above, we focus on the partial equilibrium setting in which the number of singles is kept constant. Furthermore, we first consider the special, but important, case where the reservation match qualities $\tilde{\underline{\theta}}^{mar}_{e,e'}$ and $\tilde{\underline{\theta}}^{dat}_{e,e'}$ remain unaltered with regards to changes in the material marriage benefits. In this case, we obtain the following:

Proposition 6.3 (Entry into marriage $t = 1$)**.** *Keeping the number of singles s fixed and assuming that $\frac{\partial \tilde{\underline{\theta}}^{mar}_{e,e'}}{\partial M_{\hat{e},\hat{e}'}} = \frac{\partial \tilde{\underline{\theta}}^{dat}_{e,e'}}{\partial M_{\hat{e},\hat{e}'}} = 0$, for all $e, e', \hat{e}, \hat{e}' \in E$, we have that*

$$\frac{\partial \underline{\theta}^{*}_{e,e'}}{\partial M_{e,e'}} = \frac{-1}{1 + \delta A\left(\tilde{\underline{\theta}}^{mar}_{e,e'}, \tilde{\underline{\theta}}^{dat}_{e,e'}\right)}, \tag{10}$$

where $A\left(\tilde{\underline{\theta}}^{mar}_{e,e'}, \tilde{\underline{\theta}}^{dat}_{e,e'}\right) > 0$ for κ sufficiently small.

Proof. See appendix. □

The content of Proposition 6.3 shows that, from a partial equilibrium perspective, a permanent decrease in material marriage utility $M_{e,e'}$ will make individuals in such matches pickier regarding the critical match quality required to enter marriage immediately. In other words, more people will hold off on entering marriage immediately and shift to dating as a substitute. Furthermore, Proposition 6.3 suggests there is a selection effect in marriage, where matches with a higher material surplus are more likely to immediately enter marriage, while couples with lower material benefits from marriage are more likely to use dating (cohabitation) as a substitute. Now, assuming that couples with higher educational attainment enjoy a larger material surplus from marriage (which can be rationalized by a higher level of resources for intra-household insurance), then the introduction of public insurance, though not directly targeting more highly educated individuals, mostly affects the latter. This is for two reasons: First, it is more likely to affect M for these couples (given the higher baseline surplus) and second, it is mostly the highly educated individuals who enter marriage directly instead of initially dating.

6.3.2 General equilibrium effects

The result in Proposition 6.3 is already suggestive for the sort of direct effects that result from an increase in public insurance. However, those effects only capture partial equilibrium effects in the sense that they do not control for the subsequent changes in the pool of singles or the general equilibrium effects that can imply changes in the reservation match qualities of certain match types, even when they are less or not at all directly affected by the policy change. Given the complexity of obtaining clear analytical comparative static results when taking into account the general equilibrium effects, we resort to a quantitative simulation exercise to shed light on the latter. To progress, we need to calibrate the parameters underlying our model, most importantly, the direct utility benefits from marriage by types of matches, $M_{e,e'}$. To do this, we calibrate an infinite horizon search model, calibrating the utilities from marriage by matching the implied durations of marriages by educational types, the average transition duration from singlehood into marriage and the conditional probabilities of marriages by educational attainment.[11] The result of this calibration

[11] This approach is similar to the identification strategy for an infinite horizon search model of the marriage market in Shin (2015), who studies interracial marriages. Refer to Section E in the online Appendix for full details.

exercise gives us the following values for the direct utility gains[12]:

Table 4: Material benefits from marriage

$M_{e,e'}$		e' L	e' H
e	L	0.06	0.12
e	H	0.11	0.19

Source: Authors' calculations from calibrated model. See Appendix E for full details on the calibration.

As can be seen in the table, the material benefits of marriage are much larger for couples in which both spouses have higher educational attainment than when both have lower educational attainment., i.e., $M_{H,H} > M_{L,L}$. [13] Using these marital utility flows, we simulate the model and solve for the reservation match qualities in both period $t = 1$ and $t = 2$, $\underline{\theta}^*_{e,e'}$, $\underline{\tilde{\theta}}^{mar,*}_{e,e'}$, and $\underline{\tilde{\theta}}^{dat,*}_{e,e'}$. The results are as follows:

[12]We also note that, although we can guarantee existence of a marriage market equilibrium thanks to Proposition 6.2, we cannot guarantee uniqueness, that is, a one-to-one mapping from model parameters to model outcomes. However, our calibration technique, based on Shin (2015), is robust to a multiplicity of equilibria in the sense that it involves a one-to-one map between the model outcomes suggested by the data and the model's primitives. This robustness equally carries over to the subsequent comparative statics of our calibrated model.

[13]In addition, the matrix satisfies supermodularity, in the sense that $M_{H,H} - M_{L,H} > M_{H,L} - M_{L,L}$, which implies that in a frictionless version of the model, this market would satisfy positive assortative matching, that is, material surplus would be maximized when all similar educational attainment type individuals form matches together, $M_{H,H} + M_{L,L} > M_{H,L} + M_{L,H}$. We note that this sorting result does not necessarily carry over to our setting due to the additional presence of search frictions.

Table 5: Reservation match qualities

e, e'	LL	LH	HL	HH
$\underline{\theta}^{*}_{e,e'}$	-0.05	-0.11	-0.10	-0.18
$\tilde{\underline{\theta}}^{mar,*}_{e,e'}$	0.13	-0.02	0	-0.16
$\tilde{\underline{\theta}}^{dat,*}_{e,e'}$	0.15	0	0.02	-0.14

Source: Authors' calculations from calibrated model. See Appendix E for full details on the calibration.

Table 5 confirms the intuitions we discussed in the presentation of the model. Considering the first period reservation match qualities, we find that lower-type individuals are pickier when they meet one another, compared to marriages containing at least one high-type individual. This is clearly due to the fact that the material benefits of the former type of matches ($M_{L,L}$) are lower, thereby making them less likely to simply accept any match. Furthermore, looking at the second period reservation match qualities, we find that individuals, conditional on their type of match, are pickier after dating compared to those married in the first period. This is the result of the transaction cost κ that married individuals need to pay when they decide to divorce, which acts as a barrier to separation, thereby allowing lower quality matches (in terms of match quality) to be viable. The same pattern holds for period $t = 2$ reservation match qualities as seen in the initial period: Lower-type individuals require a higher critical match quality than higher-types partnered with each other.

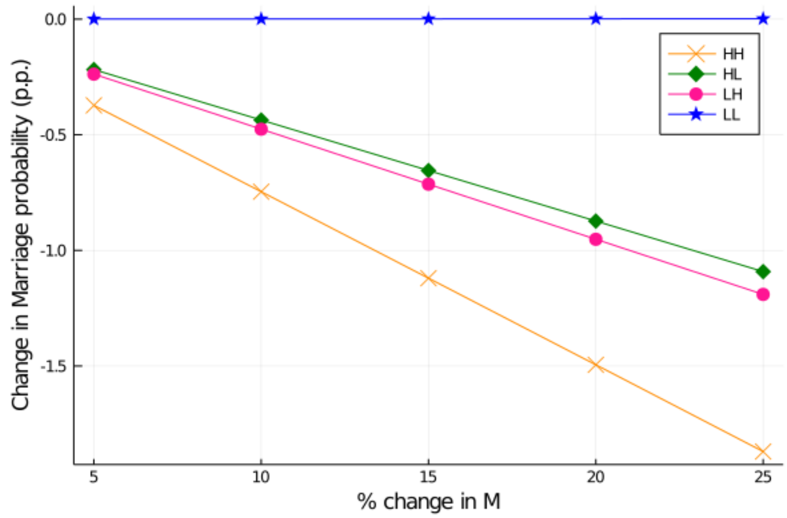
We now simulate the introduction (or expansion) of public insurance. Similar to the discussion above, we model public insurance as a reduction in the direct benefits of marriage, capturing the idea that public insurance crowds out intra-household insurance, thereby reducing the relative value of marriage. We study a scenario examining the material benefits for mixed-type and two-high-type marriages. The main reason we focus on this scenario is that the baseline material benefits of marriage are not large for matches consisting solely of less educated individuals, which suggests that public insurance is unlikely to strongly affect those matches.[14]

The results are presented in Figure 7. We experimented with several sizes for the reduction in material marriage benefits, which are plotted on the horizontal axes. We then plot the change in the entry rate into marriage in the first period, $\Delta\pi_{e,e'}$ for all $e, e' \in E$. Interestingly, we find that the decline in the entrance rate into marriage is monotonic with respect to the baseline material surplus, which confirms our hypothesis. Because of the positive relation between educational attainment and baseline material surplus (see Table 4), couples consisting of two more highly educated individuals show the greatest response in terms of delaying marriage, while couples consisting of two less-educated individuals are almost completely unaffected in terms of marriage rates. Given the positive sorting patterns typically found in the data, this result then implies a stronger responsiveness among more highly educated individuals, while less educated individuals do not respond.[15]

[14]This is also consistent with the fact that for less educated women, there seems to be a marriage health insurance penalty, as illustrated in Panel (b) of Figure 2.

[15]We further want to point out that our model yields additional implications. In particular, due to the persistence of match quality, the increased average match quality at the start of marriage will imply more stable marriages over time. However, our data are not optimal for testing this hypothesis.

Figure 7: General equilibrium effects



Source: Authors' calculations from calibrated model. Δ marriage rate for a reduction in the material benefits of both mixed matches ($\Delta M_{e,e'} < 0$, with $e \neq e'$) and HH-matches ($\Delta M_{H,H} < 0$.)

7 Conclusion

We have illustrated that public insurance can have major and sometimes counterintuitive impacts on marital outcomes using the context of the recent Medicaid expansions under the ACA. We have shown that an increase in the probability of Medicaid eligibility leads to a decrease in the marriage rate, including among those with higher levels of education. An important contribution of our work is thus to show that expansions of public insurance programs mainly targeted at lower-income individuals may still affect others in the population, in contrast to the political narrative in which public insurance programs are an important factor explaining socioeconomic gaps in marriage rates.

Our findings suggest several interesting possibilities for future research. One of the hypotheses discussed in the paper is that higher economic returns to marriage makes people less picky about match quality. It would be interesting to explore this hypotheses in different contexts and, in particular, to explore the extent to which it might explain differences in marriage and divorce rates throughout the population. Additionally, given that public insurance (partially) crowds out private insurance within marriages, there is an obvious question regarding the optimal degree of public insurance. This broad question has recently been studied in the context of unemployment insurance design (Choi and Valladares-Esteban, 2020). A similar exercise for our setting could be an interesting new application. While we focused this paper on transitions into and out of marriage, it would be interesting to study potential compositional or sorting changes in the marriage market, especially given the widespread effects we observe across the socioeconomic spectrum.

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For Online publication

A Financial Impacts of Medicaid and Health Insurance Coverage

Research has demonstrated that being uninsured in the U.S. has important negative consequences for personal financial well-being. In turn, coverage or eligibility for Medicaid has meaningful benefits for financial outcomes.

Being uninsured exposes individuals to potentially huge medical expenditures, particularly in the case of unavoidable emergency medical expenses. Health care providers typically charge uninsured patients 2-4 times more than the prices agreed upon with insurers and public programs (Garfield et al., 2019). Uninsured individuals have medical expenditures that are one half to one third the level of insured individuals for all categories of medical services except emergency medical services, for which the levels are similar (of Medicine Committee on the Consequences of Uninsurance, 2003), highlighting the fact that the uninsured are limited in their ability to control their exposure to medical emergencies.

Because of U.S. health care providers' complicated and opaque pricing systems, it's difficult to find representative statistics about the costs of medical emergencies while uninsured. The media is replete with examples of hospital stays costing thousands of dollars per night, ambulance rides that typically range from $500-1500, and even insured individuals being charged thousands and tens of thousands of dollars in out-of-pocket costs for relatively routine emergency services (see for example Sarah Kliff's extensive project on emergency room bills at Vox.) The uninsured are without a doubt vulnerable to incurring financially devastating medical expenditures in the case of emergencies. Awareness of this vulnerability is reflected in the fact that uninsured individuals in the US report much higher levels of worry about being able to pay prospective medical bills than insured individuals (Garfield et al., 2019).

Research on the effects of Medicaid on personal financial well-being has repeatedly demonstrated numerous and large benefits in a variety of settings, examined both in relation to national and local changes in policy. Medicaid coverage leads to reductions in the incidence and levels of payday loans (Fitzpatrick and Fitzpatrick, 2018; Allen et al., 2017); improvements in credit scores, reductions in debt, both medical and otherwise, and particularly in overdue debt and debt under collections (Caswell and Waidmann, 2019; Miller et al., 2018; Mazumder and Miller, 2014); and reductions in personal bankruptcy filings (Caswell and Waidmann, 2019; Gross and Notowidigdo, 2011). Conversely, researchers have also shown that the *loss* of Medicaid coverage, as in the case of Tennessee's sudden disenrollment of many individuals, led to decreases in credit scores and increases in the levels and shares of delinquent debt (Argys et al., 2017).

B Identification: discussion

We discuss more in detail the identifying assumptions needed for our analysis. A crucial ingredient in our analysis is to construct a proxy for the counterfactual measure of the likelihood to be eligible for Medicaid. To that end, suppose first that an individual's income process as a single w_i^s can be described as a function of three sources: aggregate changes in the wage returns for observable characteristics (like education or age), local labor market conditions, and idiosyncratic individual factors. Individuals' observable characteristics are denoted by a and their local labor market by l. The income process can be written as follows:

$$w_i^s = \rho_a + \rho_l + \rho_i$$

The probability that an individual would become eligible for Medicaid is then $D_i = F(w_i^s \leq m)$, where $F(w_i)$ is the cumulative distribution function for income for individual i, and m represents the applicable Medicaid eligibility threshold. In words, D_i represents the probability that an individual's income when single would fall below the Medicaid threshold. Note that although we observe realized income for currently single individuals, and can therefore exactly determine whether they are or are not eligible at a given time, D_i emphasizes the continuous probability to gain (or lose) eligibility as an ongoing process.

In practice, neither we as the econometricians nor individuals making decisions can perfectly observe D_i. However, we do observe the earnings distribution for individuals with characteristics a and individuals in labor market l. Can we use these as proxies? A concern with using earnings at the local level is that local economic conditions may be related to marital decisions, whereas we are only interested in the effect of the policy changes m. Thus, we instead use the earnings distribution for individuals with characteristics a across all locations to calculate the fraction within this group that fall below a given eligibility threshold, which we call $D_{a(i)}$.

An advantage of this assumption is that it provides a sensible approximation for one of the main challenges of causal inference with continuous treatments and observational data: estimating the probability that a particular individual would receive a particular treatment level D_i conditional on observable characteristics. Grouped by observable characteristics, $D_{a(i)}$ is the same for all individuals with characteristics a for any given level of m.

With a continuous treatment, treatment effects are described by a dose-response function, where Y_i gives a potential outcome as a function of the treatment intensity $d_{a(i)}$, for $m^* \neq m$:

$$\Delta_i = Y_i(D_{a(i)}(m^*)) - Y_i(D_{a(i)}(m)) \tag{11}$$

We consider certain restrictions on heterogeneity in treatment effects. In particular, we assume:

Assumption A (within-group treatment effects). $\Delta_{a(i)} = \Delta_i$

Assumption B (linearity)**.** *The dose-response function for each group a is linear.*

In sum, Assumption A imposes that treatment effects can differ for each group a, but not within groups, so $\Delta_{a(i)} = \Delta_i$. Assumption B further assumes that the marginal effect across the distribution of treatment D_i is constant, which we will call δ_a. It is possible that this linearity assumption may mis-specify the functional form of the relationship, but it is equally difficult to justify a specific nonlinear functional form.

Finally, we assume:

Assumption C (independence)**.** $m \perp\!\!\!\perp (Y_i)$.

In practice, Assumption C assumes that state policy changes to the eligibility threshold are exogenous. We will consider two primary ways in which this assumption might be violated through the presence of confounding relationships: first, that individuals may selectively move to states with higher eligibility, thereby generating a relationship between marital outcomes and eligibility threshold if movers have systematically different marital outcomes; and second, that states change policy in response to local economic or political conditions that may be correlated with marital outcomes.

The population average treatment effect is the expectation of δ_a:

$$ATE = \mathbb{E}[\delta_a]$$

Given the assumptions described above, identification of δ_a follows a similar logic to the case of binary treatment. Since δ_a is constant across the distribution of $D_{a(i)}$, we can write it as follows, where M reflects the actual/assigned treatment an individual experiences:

$$\mathbb{E}[\delta_a] = \mathbb{E}\left[\frac{Y_i(D_i(m^*)|M = m^*, a) - Y_i(D_i(m)|M = m^*, a)}{D(m^*) - D(m)}\right] \tag{12}$$

The fundamental problem of causal inference is clearly visible here in that the latter term in the numerator of (12) is an unobserved counterfactual. The assumption that the probability of eligibility $(D_{a(i)}(m)|M = m, a) = (D_{a(i)}(m)|M = m^*, a)$ combined with the independence of Y_i and m imply:

$$\mathbb{E}[Y_i(D_i(m)|M = m^*, a)] = \mathbb{E}[Y_i(D_i(m)|M = m, a)] \tag{13}$$

Since the latter term is observed, it can be substituted into Eq. 12, and δ_a is therefore identified.

In addition to the overall average treatment effect, we are also interested in average treatment effects by age group, level of educational attainment, and years since marriage (with divorce as the outcome), which are similarly identified as averages of δ_a for specific population subgroups.

C Medicaid State Eligibility Rules

To construct eligibility thresholds for Medicaid, we used information from annual reports of the Kaiser Commission on Medicaid and the Uninsured. These reports were based on annual surveys of state officials. In particular, we focused on the tables that indicated the income thresholds for working adults as a percent of the federal poverty line for Medicaid or Medicaid-equivalent coverage. The references below indicate which tables were used as sources in each report.

Prior to 2011, no states offered Medicaid-equivalent coverage to childless adults excepting pregnant women and individuals with certain disabilities, (although a number of states offered more limited coverage). After passage of the Affordable Care Act in 2010, states were able to request waivers for early expansion in anticipation of full expansion in 2014. As such, the reports only document eligibility for non-disabled adults beginning from reference year 2011.

In most years, the reports are published in January, documenting eligibility rules as of the January 1 of the year published. The rules are thus generally applied to eligibility for the year of publication. There are two exceptions: in 2009, a second report was published in December 2009, documenting rules in effect "as of December 2009." These rules are used for the reference year 2010, since the subsequent report was published in January 2011. Likewise, in 2013, reports were published in both January and November, and the November report included prospective eligibility rules beginning in January 2014 based on what states had announced as of October 2013. For eligibility in 2014, information from this report published in November 2013 is used.

Some states made eligibility changes midway through a given year. Since the sample we use is aggregated by annual calendar years and to maintain consistency, we use the January cutoff for changes. Thus, changes made after January are reflected only in the following year.

In certain states and years, enrollment freezes occurred that capped enrollment either at a lower level than the official threshold or stopped enrollment entirely. In these instances, we code the eligibility threshold to be either the capped lower level or zero, to capture the eligibility for enrollment in practice.

Eligibility by year is drawn from the following sources:

- Reference year 2010:
  - Cohen Ross, Donna; Jarlenski Marian; Artiga Samantha; Marks, Caryn. A foundation for health reform: Findings of a 50 State Survey of Eligibility Rules, Enrollment and Renewal Procedures, and Cost-Sharing Practices in Medicaid and Chip for Children and Parents during 2009. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Center on Budget and Policy Priorities.* December 2009. Table 3, p. 32-33.

- Reference year 2011:
  - Heberlein, Martha; Brooks, Tricia; Guyer, Jocelyn; Artiga, Samantha; Stephens, Jessica. Holding steady, looking ahead: Annual findings of a 50-State Survey of Eligibility Rules, Enrollment, and Renewal Procedures, and Cost Sharing Practices in Medicaid and Chip. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* January 2011. Table 5, p. 41-43.
- Reference year 2012:
  - Heberlein, Martha; Brooks, Tricia; Guyer, Jocelyn; Artiga, Samantha; Stephens, Jessica. Performing Under Pressure: Annual findings of a 50-State Survey of Eligibility Rules, Enrollment, and Renewal Procedures, and Cost Sharing Practices in Medicaid and Chip. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.*January 2012. Table 5, p. 42-44.
- Reference year 2013:
  - Heberlein, Martha; Brooks, Tricia; Alker, Joan; Stephens, Jessica. Getting into Gear for 2014: Findingsof a 50-State Survey of Eligibility, Enrollment, and Renewal, and Cost Sharing Practices in Medicaid and Chip. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* January 2013. Table 4, p. 33-35.
- Reference year 2014:
  - Heberlein, Martha; Brooks, Tricia; Artiga, Samantha; Stephens, Jessica. Getting into Gear for 2014: Shifting New Medicaid Eligibility and Enrollment Policies into Drive. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* November 2013. Appendix Table 2, p. 24-25.
- Reference year 2015:
  - Brooks, Tricia; Touschner, Joe; Artiga, Samantha; Stephens, Jessica; Gates, Alexandra. Modern Era Medicaid: Findings from a 50-State Survey of Eligibility, Enrollment, Renewal, and Cost-Sharing Policies in Medicaid and Chip as of January 2015. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* January 2015. Table 1, p. 24-25.

- Reference year 2016:
  - Brooks, Tricia; Miskell, Sean; Artiga, Samantha; Cornachione, Elizabeth; Gates, Alexandra. Medicaid and Chip Eligibility, Enrollment, Renewal, and Cost-Sharing Policies as of January 2016: Findings from a 50-State Survey. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* January 2016. Table 5, p. 35-37.
- Reference year 2017:
  - Brooks, Tricia; Wagnerman, Karina; Artiga, Samantha; Cornachione, Elizabeth; Ubri, Petry. Medicaid and Chip Eligibility, Enrollment, Renewal, and Cost Sharing Policies as of January 2017: Findings from a 50-State Survey. *Kaiser Commission on Medicaid and the Uninsured*, The Henry J. Kaiser Family Foundation; and *Georgetown University Center for Children and Families.* January 2017. Table 5, p. 31-32.
- Reference years 2018/2019:
  - Gifford, Kathleen; Ellis, Eileen; Coulter Edwards; Barbara, Lashbrook, Aimee; Hinton, Elizabeth; Antonisse, Larisa; Rudowitz, Robin. States focus on quality and outcomes amid waiver changes: Results from a 50-State Medicaid Budget Survey for State Fiscal Years 2018 and 2019.The Henry J. Kaiser Family Foundation, and Health Management Associates. October 2018. Table 2, p. 13

D Continuation values and Proofs

D.1 Singles' continuation values

Consider a single man with educational attainment $e \in E$, his continuation value then reads:

$$\mathcal{W}_{e,0} = 0.5 \left[\alpha_{e,H}\hat{\pi}_{e,H}\left(M_{e,H} + \mathbb{E}\left[\theta|\theta \geq -M_{e,H}\right]\right) + \alpha_{e,L}\hat{\pi}_{e,L}\left(M_{e,L} + \mathbb{E}\left[\theta|\theta \geq -M_{e,L}\right]\right)\right]$$

and similarly for a woman with educational attainment e':

$$\mathcal{W}_{0,e'} = 0.5 \left[\alpha_{H,e'}\hat{\pi}_{H,e'}\left(M_{H,e'} + \mathbb{E}\left[\theta|\theta \geq -M_{H,e'}\right]\right) + \alpha_{L,e'}\hat{\pi}_{L,e'}\left(M_{L,e'} + \mathbb{E}\left[\theta|\theta \geq -M_{L,e'}\right]\right)\right].$$

D.2 Proof of Proposition 6.2

In a first step, we focus on the reservation match qualities by types of matches and by marital status across the two periods. To that end, let us first fix the number of singles, $\left(s^m, s^f\right)$. Note that, for a given number of singles, the meeting probabilities, $\alpha_{e,e'}$ are also given. It immediately follows from (2) and (4) (which both hold at equality) that the reservation match qualities $\underline{\tilde{\theta}}_{e,e'}^{mar}, \underline{\tilde{\theta}}_{e,e'}^{dat}$ are uniquely determined given that $W_{e,0}$ and $W_{0,e'}$ are fixed. Turning to period $t = 1$, the reservation match qualities, $\underline{\theta}_{e,e'}$ are defined through (7), which should hold with equality. By using the expressions in (5) and (6), the condition for the reservation match qualities is the following:

$$M_{e,e'} + \theta + \delta\left[\varphi_{e,e'}^{mar} - \varphi_{e,e'}^{dat}\right] = 0, \tag{14}$$

where $\varphi_{e,e'}^{mar} = \mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}_{e,e'}^{mar}, 0\right\}\right]$ and $\varphi_{e,e'}^{dat} = \mathbb{E}\left[\max\left\{\tilde{\mathcal{S}}_{e,e'}^{dat}, 0\right\}\right]$. A solution to (14) is then equivalent to finding a fixed point,

$$\theta = B\left(\theta\right), \tag{15}$$

where $B\left(\theta\right) = -M_{e,e'} - \delta\left[\varphi_{e,e'}^{mar} - \varphi_{e,e'}^{dat}\right]$. Note that, given the number of singles is fixed, we can treat each equation in this system of equations separately. To proceed, we define $C_{e,e'}^{mar} = -\underline{\tilde{\theta}}_{e,e'} + \mathbb{E}\left[\tilde{\theta}|\tilde{\theta} \geq \underline{\tilde{\theta}}_{e,e'}\right]$ (and a similar expression for $C_{e,e'}^{dat}$). We then have the following:

Lemma D.1. *Given the number of singles, we have*

$$\frac{\partial\left(\varphi_{e,e'}^{mar} - \varphi_{e,e'}^{dat}\right)}{\partial\theta} = A\left(\underline{\tilde{\theta}}_{e,e'}^{mar}, \underline{\tilde{\theta}}_{e,e'}^{dat}\right), \tag{16}$$

where

$$\begin{aligned} A\left(\underline{\tilde{\theta}}_{e,e'}^{mar}, \underline{\tilde{\theta}}_{e,e'}^{dat}\right) = \pi_{e,e'}^{mar}\ \mathbb{E}\left[\tilde{\theta}\left(\tilde{\theta} - \theta\right) \mid \underline{\tilde{\theta}}_{e,e'}^{mar} \leq \tilde{\theta} \leq \underline{\tilde{\theta}}_{e,e'}^{dat}\right] - \left(\pi_{e,e'}^{dat} - \pi_{e,e'}^{mar}\right)\mathbb{E}\left[\tilde{\theta}\left(\tilde{\theta} - \theta\right) \mid \tilde{\theta} \geq \underline{\tilde{\theta}}_{e,e'}^{dat}\right] \\ C_{e,e'}^{mar}\ \mathbb{E}\left[\left(\tilde{\theta} - \theta\right) \mid \underline{\tilde{\theta}}_{e,e'}^{mar} \leq \tilde{\theta} \leq \underline{\tilde{\theta}}_{e,e'}^{dat}\right] - \left(C_{e,e'}^{dat} - C_{e,e'}^{mar}\right)\ \mathbb{E}\left[\left(\tilde{\theta} - \theta\right) | \tilde{\theta} \geq \underline{\tilde{\theta}}_{e,e'}^{dat}\right]. \end{aligned} \tag{17}$$

Proof. First note that $\varphi_{e,e'}^{mar} = \pi_{e,e'}^{mar} C_{e,e'}^{mar}$, which implies that

$$\frac{\partial\varphi_{e,e'}^{mar}}{\partial\underline{\theta}_{e,e'}} = \frac{\partial\pi_{e,e'}^{mar}}{\partial\theta} C_{e,e'}^{mar} + \pi_{e,e'}^{mar}\frac{\partial C_{e,e'}^{mar}}{\partial\theta}.$$

We have that:

$$\frac{\partial\pi_{e,e'}^{mar}}{\partial\theta} = \int_{\underline{\tilde{\theta}}_{e,e'}^{mar}}^{+\infty}\left(\tilde{\theta} - \theta\right) g\left(\tilde{\theta}|\theta\right) d\tilde{\theta}, \tag{18}$$

and

$$\frac{\partial C_{e,e'}^{mar}}{\partial\theta} = \int_{\underline{\tilde{\theta}}_{e,e'}^{mar}}^{+\infty}\tilde{\theta}\left(\tilde{\theta} - \theta\right) g\left(\tilde{\theta}|\theta\right) d\tilde{\theta}. \tag{19}$$

Obviously, similar expressions hold for $\frac{\partial \pi^{dat}_{e,e'}}{\partial \theta}$ and $\frac{\partial C^{dat}_{e,e'}}{\partial \theta}$. We then have that:

$$\begin{aligned}\frac{\partial \varphi^{mar}_{e,e'}}{\partial \theta} - \frac{\partial \varphi^{dat}_{e,e'}}{\partial \theta} =& C^{mar}_{e,e'}\ \mathbb{E}\left[\left(\tilde{\theta}-\theta\right) \mid \tilde{\theta} \geq \underline{\tilde{\theta}}^{mar}_{e,e'}\right] - C^{dat}_{e,e'}\ \mathbb{E}\left[\left(\tilde{\theta}-\theta\right) \mid \tilde{\theta} \geq \underline{\tilde{\theta}}^{dat}_{e,e'}\right] \\ &+ \pi^{mar}_{e,e'}\ \mathbb{E}\left[\tilde{\theta}\left(\tilde{\theta}-\theta\right) \mid \tilde{\theta} \geq \underline{\tilde{\theta}}^{mar}_{e,e'}\right] - \pi^{dat}_{e,e'}\ \mathbb{E}\left[\tilde{\theta}\left(\tilde{\theta}-\theta\right) \mid \tilde{\theta} \geq \underline{\tilde{\theta}}^{dat}_{e,e'}\right]\end{aligned} \tag{20}$$

After some basic algebra, we obtain

$$\frac{\partial\left(\varphi^{mar}_{e,e'} - \varphi^{dat}_{e,e'}\right)}{\partial \theta} = A\left(\underline{\tilde{\theta}}^{mar}_{e,e'}, \underline{\tilde{\theta}}^{dat}_{e,e'}\right),$$

where A is given as in (17), which is the desired result. □

Notice that, when $\kappa \to 0$, $\underline{\tilde{\theta}}^{mar}_{e,e'} - \underline{\tilde{\theta}}^{dat}_{e,e'} \to 0$ for all $e, e' \in E$, which means that there exists a $\epsilon > 0$ such that, if $|\underline{\tilde{\theta}}^{mar}_{e,e'} - \underline{\tilde{\theta}}^{dat}_{e,e'}| < \epsilon$ then $|A\left(\underline{\tilde{\theta}}^{mar}_{e,e'}, \underline{\tilde{\theta}}^{dat}_{e,e'}\right)| < 1$. As a consequence of Lemma D.1,

$$B'(\theta) = -\delta\frac{\partial\left(\varphi^{mar}_{e,e'} - \varphi^{dat}_{e,e'}\right)}{\partial \theta} = -\delta A\left(\underline{\tilde{\theta}}^{mar}_{e,e'}, \underline{\tilde{\theta}}^{dat}_{e,e'}\right).$$

Given that $\delta < 1$, and $|A\left(\underline{\tilde{\theta}}^{mar}_{e,e'}, \underline{\tilde{\theta}}^{dat}_{e,e'}\right)| < 1$ for $\kappa < \epsilon$, we then immediately obtain that $|B'(\theta)| < 1$. We conclude that (15) has a (unique) solution, say $\underline{\theta}$ for each e, e' as a consequence of Banach's fixed point theorem. Furthermore notice we can suitably redefine the domain for the vector of reservation match qualities θ. In particular, for the period $t = 1$ reservation match qualities, we can define $\underline{\theta}^0 = B(+\infty)$, and $\underline{\theta}^1 = B\left(\underline{\theta}^0\right)$, we can then restrict $\underline{\theta} \in \left[\underline{\theta}^0, \underline{\theta}^1\right]$. Similarly, for the second period reservation match qualities we can define $\underline{\tilde{\theta}}^{mar,0}_{e,e'} = -M_{e,e'} - \kappa$ and $\underline{\tilde{\theta}}^{mar,1}_{e,e'} = -M_{e,e'} - \kappa + 0.5\sum_{e'} \gamma^m_e \gamma^f_{e'} \hat{\pi}_{e,e'}\left(M_{e,e'} + \mathbb{E}\left[\theta | \theta \geq -M_{e,e'}\right]\right)$. For the reservation match qualities after dating we can define bounds analogously, $\underline{\tilde{\theta}}^{dat,0}_{e,e'} = \underline{\tilde{\theta}}^{mar,0}_{e,e'} + \kappa$ and $\underline{\tilde{\theta}}^{dat,1}_{e,e'} = \underline{\tilde{\theta}}^{mar,1}_{e,e'} + \kappa$. So the domain can be restricted to the hypercubes $\left[\underline{\tilde{\theta}}^{mar,0}, \underline{\tilde{\theta}}^{mar,1}\right]$ and $\left[\underline{\tilde{\theta}}^{dat,0}, \underline{\tilde{\theta}}^{dat,1}\right]$. Now, note that we can write

$$s^m_e = D^m_e\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar}, \underline{\tilde{\theta}}^{dat}\right) \tag{21}$$

$$s^f_{e'} = D^f_{e'}\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar}, \underline{\tilde{\theta}}^{dat}\right), \tag{22}$$

where

$$D_e^m\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar},\underline{\tilde{\theta}}^{dat}\right)=\gamma_e^m\sum_{e'}\gamma_{e'}^f\left(\left[1-F\left(\underline{\theta}_{e,e'}\right)\right]\int_{\underline{\theta}_{e,e'}}^{+\infty}G\left(\underline{\tilde{\theta}}_{e,e'}^{mar}|\theta\right)f\left(\theta\right)d\theta\right.$$

$$\left.+F\left(\underline{\theta}_{e,e'}\right)\int_{-\infty}^{\underline{\theta}_{e,e'}}G\left(\underline{\tilde{\theta}}_{e,e'}^{dat}|\theta\right)f\left(\theta\right)d\theta\right),\text{ and}$$

$$D_{e'}^f\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar},\underline{\tilde{\theta}}^{dat}\right)=\gamma_{e'}^f\sum_{e}\gamma_{e}^m\left(\left[1-F\left(\underline{\theta}_{e,e'}\right)\right]\int_{\underline{\theta}_{e,e'}}^{+\infty}G\left(\underline{\tilde{\theta}}_{e,e'}^{mar}|\theta\right)f\left(\theta\right)d\theta\right.$$

$$\left.+F\left(\underline{\theta}_{e,e'}\right)\int_{-\infty}^{\underline{\theta}_{e,e'}}G\left(\underline{\tilde{\theta}}_{e,e'}^{mar}|\theta\right)f\left(\theta\right)d\theta\right).$$

We now consider the following

$$s_e^{m,0}=\min_{\underline{\theta}\in\left[\underline{\theta}^0,\underline{\theta}^1\right]}D_e^m\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar,0},\underline{\tilde{\theta}}^{dat,0}\right),\text{ and}$$

$$s_{e'}^{f,0}=\min_{\underline{\theta}\in\left[\underline{\theta}^0,\underline{\theta}^1\right]}D_{e'}^f\left(\underline{\theta},\ \underline{\tilde{\theta}}^{mar,0},\underline{\tilde{\theta}}^{dat,0}\right).$$

Both minimization problems have a solution given Weierstrass' extreme value theorem. Then $s^m\in\left[s^{m,0},\gamma^m\right]$ and $s^f\in\left[s^{f,0},\gamma^f\right]$. Taking stock, we have shown that there exists continuous solutions $\underline{\tilde{\theta}}_{e,e'}^{mar}\left(s^m,s^f\right),\underline{\tilde{\theta}}_{e,e'}^{dat}\left(s^m,s^f\right)$, and $\underline{\theta}_{e,e'}\left(s^m,s^f\right)$. Using these, we can then construct mappings, $D_e^g\left(\underline{\theta}\left(s^m,s^f\right),\ \underline{\tilde{\theta}}^{mar}\left(s^m,s^f\right),\underline{\tilde{\theta}}^{dat}\left(s^m,s^f\right)\right)$ for $g=m,f$ and all $e\in E$. Note that these maps are continuous in $\left(s^m,s^f\right)$. The measures of singles then have to satisfy the following:

$$s_e^m=D_e^m\left(\underline{\theta}\left(s^m,s^f\right),\ \underline{\tilde{\theta}}^{mar}\left(s^m,s^f\right),\underline{\tilde{\theta}}^{dat}\left(s^m,s^f\right)\right),\tag{23}$$

$$s_{e'}^f=D_{e'}^f\left(\underline{\theta}\left(s^m,s^f\right),\ \underline{\tilde{\theta}}^{mar}\left(s^m,s^f\right),\underline{\tilde{\theta}}^{dat}\left(s^m,s^f\right)\right).\tag{24}$$

Given that the domain for $\left(s^m,s^f\right)$ is the hypercube $\left[s^{m,0},\gamma^m\right]\times\left[s^{f,0},\gamma^f\right]$, we can immediately apply Brouwer's fixed point theorem and we know there exists a solution $\left(s^{m*},s^{f*}\right)$ to (23) and (24). We have thereby established the existence of a marriage market equilibrium.

D.3 Proof of Proposition 6.3

Similar as in the proof of Proposition 6.2, we start with the condition for the reservation match qualities in period $t=1$,

$$M_{e,e'} + \theta + \delta \left[\varphi^{mar}_{e,e'} - \varphi^{dat}_{e,e'}\right] = 0.$$

Using implicit differentiation with respect to $M_{e,e'}$, we obtain:

$$\frac{\partial \underline{\theta}^*_{e,e'}}{\partial M_{e,e'}} + \delta \frac{\partial \left(\varphi^{mar}_{e,e'} - \varphi^{dat}_{e,e'}\right)}{\partial \theta} \frac{\partial \underline{\theta}^*_{e,e'}}{\partial M_{e,e'}} = -1 - \delta \frac{\partial \left(\varphi^{mar}_{e,e'} - \varphi^{dat}_{e,e'}\right)}{\partial M_{e,e'}}. \tag{25}$$

Keeping the number of singles constant and assuming that $\frac{\partial \tilde{\underline{\theta}}^{mar}_{e,e'}}{\partial M_{\hat{e},\hat{e}'}} = \frac{\partial \tilde{\underline{\theta}}^{dat}_{e,e'}}{\partial M_{\hat{e},\hat{e}'}} = 0$, Lemma D.1 implies that (25) can be rewritten as:

$$\frac{\partial \underline{\theta}^*_{e,e'}}{\partial M_{e,e'}} + \delta A\left(\tilde{\underline{\theta}}^{mar}_{e,e'}, \tilde{\underline{\theta}}^{dat}_{e,e'}\right) \frac{\partial \underline{\theta}^*_{e,e'}}{\partial M_{e,e'}} = -1,$$

which immediately gives the desired result.

E Details on the calibration

To simulate the general equilibrium effects of public insurance, we opted to select numerical values for the (baseline) material benefits that have ties with empirical findings. To be more precise, we opted to calibrate a rather standard, continuous time and infinite horizon version of a search-based model of marriage, with some simplifications for easing the identification,

- Match qualities follow a Poisson process, with λ the arrival parameter of a change in match quality, drawn from some distribution H.
- There is no initial dating, individuals choose to (re-)marry or become single.
- Meeting is random, with a search technology, $T\left(S^m, S^f\right)$ that is increasing in both arguments. This technology determines the arrival rates of matches,

  $$\alpha^m_{e,e'} = \frac{T\left(S^m, S^f\right)}{S^m S^f} s^f_{e'}$$

  denotes the arrival rate for men with educational attainment $e \in E$ meeting women of educational attainment $e' \in E$ and similarly for women with education level e' meeting a man of type e,

  $$\alpha^f_{e,e'} = \frac{T\left(S^m, S^f\right)}{S^m S^f} s^m_{e}$$

- A discount rate r used for discounting,

In this model, it can be shown that the reservation match qualities can be obtained as the solution of the following system of equations:

$$\underline{\theta}_{e,e'} + \frac{\lambda}{r+\lambda}\varphi\left(\underline{\theta}_{e,e'}\right) + M_{e,e'} = r\mathcal{V}^m_{e,0} + r\mathcal{V}^f_{0,e'}, \tag{26}$$

$$r\mathcal{V}^m_{e,0} = \frac{1}{r+\lambda}\left[0.5\sum_{e''}\alpha^m_{e,e''}\varphi\left(\underline{\theta}_{e,e''}\right)\right], \tag{27}$$

$$r\mathcal{V}^f_{0,e'} = \frac{1}{r+\lambda}\left[0.5\sum_{e''}\alpha^f_{e,e''}\varphi\left(\underline{\theta}_{e'',e'}\right)\right]. \tag{28}$$

Where $\varphi\left(\underline{\theta}_{e,e'}\right) = \int_{\underline{\theta}_{e,e'}}^{\infty}\left[1 - H\left(\theta\right)\right]d\theta$, denotes the expected maximum surplus of a match where the reservation match quality is given by $\underline{\theta}_{e,e'}$. Condition (26) essentially states that the current spouses are indifferent between staying married (and receiving a total (expected) surplus as given by the left hand side expression) and divorcing, the latter giving each spouse their resp. continuation value of being single, in particular $\mathcal{V}^m_{e,0}$ for a male of type e and $\mathcal{V}^f_{0,e'}$ for a woman of type e'. Equations (27) and (28) define the continuation values for men and women respectively.

We are particularly interested in the following empirical objects:

- $D^1_{e,e'}$ = the average duration of marriages of type e, e',
- $D^{0,g}_e$ = the average duration of individuals of gender $g \in \{m, f\}$ and educational attainment $e \in E$ transitioning from singlehood to marriage,
- $P^m_{e,e'}$ = the probability of a male with educational attainment e to marry a woman with educational level e', conditional on marrying. A similar object can be constructed for women across educational types, $P^f_{e,e'}$.

For $D^1_{e,e'}$ and $P^m_{e,e'}$, we use data from the Panel Study on Income Dynamics. Specifically, we consider individuals in the PSID married between 1990-2000 (to allow for some meaningful variation in marriage duration). Then, we calculate the average duration of these marriages, censoring the highest value at 19 (since our data goes until 2019, this is the longest possible duration for those married in 2000). Next, we calculate the fraction of marriages over this period by type e, e' to represent the object $P^m_{e,e'}$.

For $D^{0,g}_e$, due to a small number of observations after imposing our sample restrictions and splitting by educational attainment, we use evidence presented in Aughinbaugh et al. (2013) (who use the NLSY79, a nationally representative sample of men and women in the US and who were between 14 and 22 years old at the time of their first interview in 1979) on average duration till remarriage, broken down by educational attainment. To be more precise, the average duration to remarriage from time of divorce is about 3-4 years, with a lower duration for individuals who have at least a bachelor's degree. To match these statistics, we picked an average duration of 2 years to remarry for individuals with a high level of education, and 5 for individuals with low levels of education. Given the type distribution across education, we match the aggregate averages presented in Aughinbaugh et al. (2013).

Our finalized data thus consists in

$$\left\{ \left(D^{1}_{e,e'}\right)_{e,e'\in E}, \left(D^{0,m}_{e}\right)_{e\in E}, \left(D^{0,f}_{e'}\right)_{e'\in E}, \left(P^{m}_{e,e'}\right)_{e,e'\in E}, \left(P^{f}_{e,e'}\right)_{e,e'\in E} \right\}.$$

Using this, we can then calibrate some of the structural parameters. To start, we note that,

$$\mathbb{E}\left[D^{1}_{e,e'}\right] = \frac{1}{\lambda H\left(\underline{\theta}_{e,e'}\right)}.$$

Given a calibrated value for λ, we can then immediately compute the implied rejection probability,

$$\hat{G}\left(\underline{\theta}_{e,e'}\right) = \frac{1}{\bar{D}^{1}_{e,e'}},$$

where $\bar{D}^{1}_{e,e'}$ is the average duration of marriages of type e, e'.

Similarly, note that

$$\mathbb{E}\left[D^{0,m}_{e}\right] = \frac{1}{\sum_{e''} \alpha^{m}_{e,e''}\left[1 - G\left(\underline{\theta}_{e,e''}\right)\right]},$$

and

$$P^{m}_{e,e'} = \frac{\alpha^{m}_{e,e'}\left[1 - G\left(\underline{\theta}_{e,e'}\right)\right]}{\sum_{e''} \alpha^{m}_{e,e''}\left[1 - G\left(\underline{\theta}_{e,e''}\right)\right]}.$$

Combining both yields the following:

$$\alpha^{m}_{e,e'} = \frac{P^{m}_{e,e'}}{\mathbb{E}\left[D^{0,m}_{e}\right]} \frac{1}{1 - \frac{1}{\lambda \mathbb{E}\left[D^{1}_{e,e'}\right]}}.$$

And from this, we can immediately approximate the arrival rates from the data as follows:

$$\hat{\alpha}^{m}_{e,e'} = \frac{P^{m}_{e,e'}}{\bar{D}^{0,m}_{e}} \frac{1}{1 - \frac{1}{\lambda \bar{D}^{1}_{e,e'}}}.$$

With the estimated acceptance (or equivalently, rejection) probabilities and arrival rates of marriages, we can then proceed with calibrating deeper structural objects from the model. In particular, using (26) and (27)-(28) and the calibrated values for the acceptance probabilities and arrival rates we can now compute the implied (net) utility flow of a marriage of a particular type, $M_{e,e'}$.