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

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Abstract

While many marry for love, marriage is also a legal and economic institution. Public insurance programs for individuals decrease the relative insurance value of marriage. We explore how this tradeoff affects the marriage market in the context of the Affordable Care Act's Medicaid expansion in the United States. We show both theoretically and empirically that Medicaid expansion does decrease the marriage rate – even for individuals with high educational attainment – and it also reduces the divorce rate of new marriages, consistent with an increase in match quality. Counter to intuition, there is no effect on divorce rates for already-married couples. Our findings illustrate that even when public insurance reduces the monetary benefits of marriage, the effects on marriage overall may be positive.

Keywords: Marital quality, Health insurance, Medicaid expansions, Intra-household allocation

JEL: D13, I38, J12

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1 Introduction

Marriage as an institution provides an important source of private insurance against a wide range of shocks (both at the aggregate and individual level). At the same time, most high-income countries also provide insurance to individuals collectively through public safety net programs. There have long been serious debates about whether such public programs crowd out or even disincentivize marriage by reducing its relative economic value compared to singlehood. Generally, the focus has been on the quantity of marriages. In this paper, we explore another dimension of the effects of expansions in public insurance: the quality of marriages.

This paper aims to contribute to this discussion by focusing on a particular policy change which has substantially expanded public insurance. We exploit the expansion of Medicaid in the United States following the passage of the Patient Protection and Affordable Care Act in 2010 (more commonly referred to as the ACA or "Obamacare") and study its effects on the marriage market.

Married individuals in the United States have historically benefited from more options for health insurance coverage than their non-married peers, since individuals can be covered via their spouse's (employer-sponsored) health insurance plan if they lack or lose coverage themselves. The expansion of Medicaid, a means-tested public health insurance program, provides an additional source of health insurance coverage for many unmarried adults.

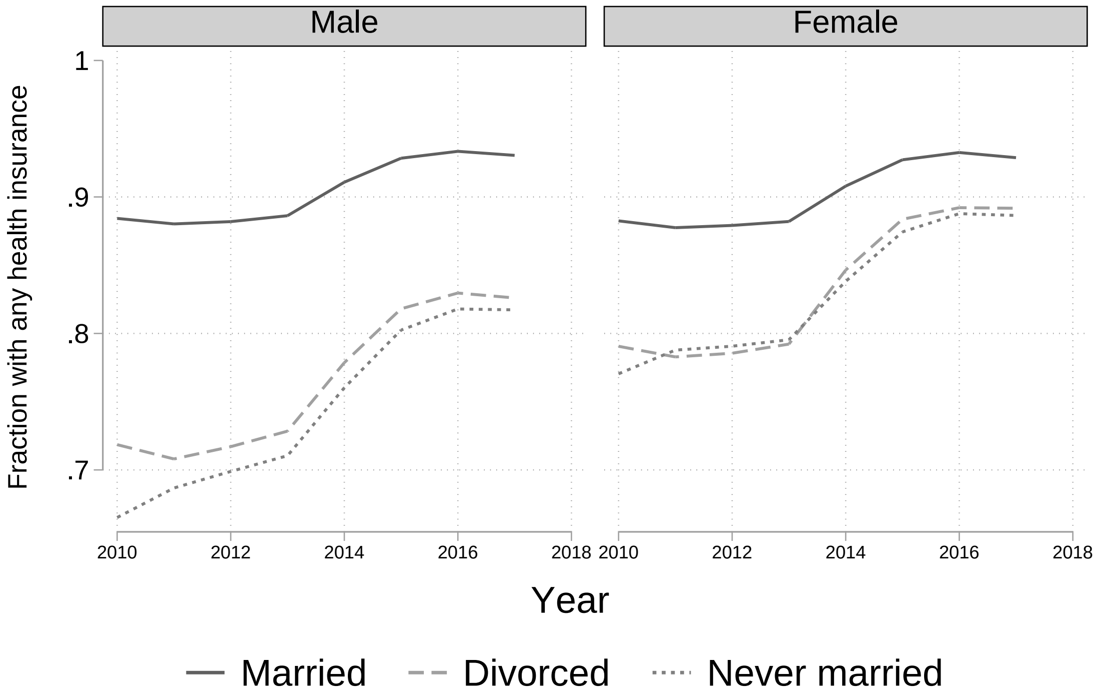
Fig. 1 highlights the increase in health insurance coverage for unmarried compared to married adults following the implementation of the ACA Medicaid expansion. The gaps in health insurance coverage between married and non-married adults shrunk markedly around the year 2014, when many key ACA reforms including Medicaid expansion were largely implemented. 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.

Furthermore (and perhaps uniquely among high-income countries), health insurance in the United States is not only important for health but also affects exposure to financial shocks. Accessing medical services (and particularly emergency medical services) while uninsured can cost thousands or tens of thousands of dollars and routinely puts households into debt and bankruptcy.[1] We posit that the Medicaid expansions should primarily affect marital outcomes via this financial channel.

What we do. Previous work studying the effects of Medicaid, even in the context of its impact on financial outcomes, has mostly focused on its direct impacts via increased coverage and eligibility. However, in relation to the marriage market, we hypothesize that there may also be substantial indirect effects of Medicaid expansion. In particular, when making marital decisions, currently married individuals may take into account the

[1] Sec. 2.3 provides a detailed review of research on the financial impacts of non-insurance and gaining/losing Medicaid.

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 2010-2017 for adults between the ages of 18 and 64.

availability of public insurance programs for non-married individuals, which thereby affects their outside options (e.g. through the likelihood to find a new partner and through being eligible for public insurance outside of marriage.)

Building on this intuition, the goal of our empirical analysis is to assess the causal effect on the marital outcomes of an increase in the probability of Medicaid coverage when not married. Then, we develop a search-and-matching model of the marriage market to provide insights on the mechanisms and interpretation of our empirical effects.

Empirically, we adopt an identification strategy similar in spirit to the simulated instruments approach (originally implemented by Currie and Gruber (1996a,b), and widely used in work related to Medicaid). We use a similarly constructed measure of the probability that non-married individuals with certain demographic characteristics would be eligible for Medicaid in a certain state and year, $PrMedicaid$, as a direct proxy for the unobserved probability to be eligible for Medicaid when non-married.

We use the American Community Survey (ACS) over the years 2010-2019 to conduct our empirical analysis. We show that an increase in $PrMedicaid$ decreases the marriage rates for young men and women in their twenties, an age range when most first marriages occur.

Additionally, a novel finding among these results is that even individuals with high levels of educational attainment, across all age groups, have decreased marriage rates in response to increased Medicaid eligibility. Typically, most work that studies responses to means-tested programs only study individuals with low income or low educational attainment, and this finding highlights that these programs may in fact have broader effects, since even individuals with higher educational attainment–and higher expected earnings–may take into account the risks of negative shocks under which they might benefit from means-tested programs.

We also find that an increase in the probability of eligibility *at the time of marriage* decreases subsequent divorce rates of newly formed married couples. This finding is suggestive that Medicaid expansion may increase positive selection into marriage.

Finally, we found little evidence of a net effect on divorce for already-married couples in response to increases in the probability of eligibility for Medicaid outside of marriage. While one might think at first glance that an improvement in the outside option for already-married individuals would lead to increases in divorce, our theoretical framework highlights the importance of considering general equilibrium effects of the increase in the effect of public insurance expansion, and implies that the direction of the net effect is in fact ambiguous.

We address several potential threats to our identification strategy. In particular, endogenous migration to states with more generous Medicaid rules as well as state-level policy endogeneity, where state decisions respond to local conditions correlated with marital outcomes, could each be possible issues. In fact, for migration we find that for men there is a small but significant increase in the rate of individuals who recently moved to their new state of residence, for states with higher Medicaid eligibility. However, migration does not appear to function as a confounder, since either excluding those who recently changed states or including a control for the fraction of recent movers in each demographic group and state leaves estimates essentially unchanged.

While in our baseline specifications, we include all Medicaid eligibility changes between 2010-2019, in robustness checks assessing the role of policy endogeneity, we adapt the approach of **?** to effectively exploit changes owing specifically to the ACA Medicaid expansion within states that did or did not have as Democratic governor as of 2011. The key assumption within this strategy is that within these groups, the decision to expand Medicaid is as-good-as-random. As for the question of migration, this robustness check has little effect on the main results.

In the second part of the paper, we formulate a stylized search-based equilibrium model of the marriage market to rationalize and aid in interpretation of these results, as well as to highlight that the same mechanisms could be at work more generally in other types of public insurance programs. The model is based on that of Anderberg (2007).

Individuals in our model face both financial and match quality shocks. They decide whether or not to enter marriage. There are two benefits to marriage: first, spouses provide insurance to each other through intra-household transfers. Second, there are immaterial benefits of marriage, which we refer to (in line with the literature) as match quality, and which one can interpret as 'love' in the relationship. Match quality evolves randomly and spouses can dissolve the marriage whenever the marital surplus is too low. Searching for a potential spouse is costly in terms of time spent to find an acceptable spouse. In the equilibrium without public insurance, each individual uses a cut-off acceptance strategy, accepting a potential spouse if the match quality is above a certain threshold level.

The inclusion of public insurance increases the value of singlehood vis a vis marriage.[2] As a result, the critical match quality level for individuals to choose marriage increases. The increase in critical match quality owes to an improvement in the outside option: public insurance is a substitute for intra-household insurance, so individuals become pickier in the marriage market. This leads to a decrease in the marriage rate.

With pickier individuals from the start, the average match quality of newly formed matches will be higher. If we then further assume that match quality features some persistence, newly formed marriages will face a lower probability of divorce. Finally, the model also predicts that intra-household allocations will shift if public insurance is more valuable to one spouse than the other.

Notwithstanding its relative simplicity, our model is informative about the mechanisms through which public insurance might affect the marriage market, both in terms of changes in the rates of marriage formation and dissolution of marriages. We also discuss how the model can incorporate richer features, e.g. the implications on our search-based equilibrium in case individuals might also care about producing and sharing public goods within the household.

Our contribution relative to existing literature. Our paper makes several contributions. First, our study of the effects of the ACA-era Medicaid expansion provides insights on the implications of expanding more general forms of public insurance on marital outcomes. While a large body of work has studied the relationship between U.S. welfare programs and marriage or family structure (e.g. Bitler et al. (2004); Moffitt (1990); Hoffman and Duncan (1995); Hoynes (1997)), including a few such papers related to Medicaid reforms (Yelowitz, 1998; Decker, 2000; Farley, 2001), the policies studied in this literature are relatively narrow and specific to the U.S. context.[3] In contrast, the broad-based expansion of Medicaid to all low-income individuals is more easily generalizable to other types of public insurance and might therefore be more informative to address the question as to how public insurance affects the marriage market and intra-household risk insurance.

Second, we use insights from a general equilibrium model of the marriage market to demonstrate that Medicaid (and public insurance) can affect not only marriage and divorce probabilities, but also match quality and intra-household allocations. These insights build further upon the theoretical model of Anderberg (2007), who analyzes the relationship between intra-household consumption risk-sharing and public insurance on marital outcomes.

Third, our contribution is most similar to earlier work showing that public policy

[2]When transfers are means-tested with eligibility thresholds depending on household size, as is the case with Medicaid, the expected value of public insurance is higher for singles than for married households, because the likelihood to become eligible is higher.

[3]For example, two out of the three papers cited on Medicaid policy and marriage are about the effect of Medicaid asset eligibility limits in causing "medical" divorce when one spouse is ill.

changes which affect the value of marriage relative to singlehood can impact the marriage market equilibrium and intra-household allocations (Chiappori and Oreffice, 2008; Low et al., 2018) as well as match quality (Persson, 2020). Notably, Chiappori and Oreffice (2008) and Persson (2020) study implications of a universal policy change that applies to the whole population, and Low et al. (2018) study the effects of a change to a means-tested program, but with a focus on its impact on women with low-educational attainment. In contrast, we study the effects of expansion of a means-tested program on the whole population.

It is not self-evident that a means-tested program would have effects for individuals across the population. However, this turns out to be a final novel contribution of our work, which is to highlight that even policy targeted towards those with lower-incomes can have effects on marriage quality and intra-household allocations broadly across the population. Our findings suggest that public safety net programs like Medicaid may impact the behavior of individuals, even those individuals who have a low propensity to currently (or ever) becoming eligible for coverage, implying broader impacts than previously recognized.

Overview for the rest of the paper. The structure for the rest of the paper is as follows: Section 2 provides the institutional background on the Affordable Care Act and in particular the important components of the Medicaid expansion. A discussion of the dataset we will use, our identification strategy and the empirical predictions are contained in section 3. Section 4 contains the empirical results and the robustness checks pertaining to our empirical analysis. Section 6 concludes the paper. Section 5 presents the search-based model of the marriage market which we will use as a theoretical framework to understand the direct and indirect implications of the Medicaid expansion on the marriage market.

2 Institutional Background

In 2010, the U.S. Congress passed the Patient Protection and Affordable Care Act, a sweeping reform colloquially known as the ACA or Obamacare. The ACA changed the options for health insurance coverage for individuals in a number of important ways compared to the previously existing options.

2.1 Prior to the ACA

About three-fifths of the non-elderly population in the United States were covered by employer-sponsored insurance (ESI) prior to the Affordable Care Act's passage (Long et al., 2016). ESI is a fringe benefit tied to an individual's job. Spouses can typically be covered on the employee's plan for an additional premium. Thus, non-working spouses or those who have lost their own ESI can still access coverage by virtue of marriage.

Beyond ESI, there were a few other sources of coverage. The elderly (above the age of 64) were (and still are) covered by the universal public program Medicare (they are thus excluded from our analysis). Individuals could also purchase private non-group insurance, however only about 6% of the non-elderly population was covered in this way prior to the reform.

A final major source of coverage for the non-elderly was Medicaid, the federal name for state-level public health insurance programs intended to cover particularly disadvantaged low-income groups. Though states had substantial flexibility in administering their programs, by the time of the ACA there were federal requirements to provide coverage for low-income pregnant women and children, parents who met requirements for state cash welfare programs (formerly AFDC) and disabled individuals who qualified for Supplemental Security Income. Additionally, many states throughout the 1990s and early 2000s expanded eligibility for parents further, leading to substantial variation in eligibility thresholds for working parents: by 2012, 17 states restricted parental eligibility to levels under 50% of the federal poverty line, while 18 states had eligibility levels over 100% of the federal poverty line (Musumeci, 2012).

In the year prior to the passage of the ACA the non-elderly adult uninsured rate was about 20%. With the structure of ESI offers and Medicaid eligibility rules, the uninsurance rate was markedly higher for lower-income individuals: between 25-30% of those under 400% of the FPL were uninsured, compared to only 14% of those above 400% of the FPL.

2.2 ACA Medicaid Expansion

Medicaid expansion was one among several measures in the ACA intended to reduce uninsurance. As written, the legislation mandated the expansion of Medicaid to cover all individuals under 138% of the federal poverty line (Foundation, 2010). The Medicaid expansion was intended to be implemented in 2014 nationally, with the possibility for state waivers to begin expansion earlier. Those with offers of employer-sponsored insurance were not to be eligible for either Medicaid or premium tax subsidies for the purchase of private insurance, another of the key ACA initiatives.

Upon passage of the ACA, 25 states sued the federal government in opposition to the Medicaid expansion. In 2012 the Supreme Court ruled that the federal government could not condition states' other sources of federal funding on participation in the expansion. As a result, participation in the Medicaid expansion effectively became optional (Musumeci, 2012).

Six states took up the option for early, partial expansion of Medicaid by 2012 (Foundation, 2012), and a further 19 states expanded according to the originally planned ACA roll-out in 2014. A handful of additional states have expanded Medicaid in the years since, bringing the total to 37 states as of the beginning of 2020.

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 both categories in eligibility rules, in particular by parental status. We exploit this additional variation in our analysis.[4]

[4] We provide more detail on this point in App. B, where we describe the sources used to code eligibility rules across states and individual characteristics.

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

2.3 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 aggregate 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.) It seems safe to say that the uninsured are 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).

Perhaps not surprisingly, 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).

Given this body of research, we are interested in exploring the impact of Medicaid on marriage via the insurance Medicaid provides against negative income shocks.

3 Empirical Strategy

The goal of the empirical strategy is to estimate average causal effects associated with an increase in the probability of becoming eligible for Medicaid when non-married on marital outcomes. We use a strategy that is inspired by the simulated instruments approach pioneered by Currie and Gruber (1996a), which constructs an estimate of the likelihood that individuals will be eligible for a program. In this section, we discuss assumptions necessary for identification, possible challenges for these assumptions, the data we use for estimation, and details of estimation given our data, identification assumptions, and estimands of interest.

3.1 Identification

Unlike many settings that study health insurance, we are not interested solely in the effects of actual coverage or eligibility for coverage (although these may play a role). This is a typical application of simulated instruments, where the measure serves as an instrument for the binary treatment of interest. In our setting, the same measure used in other papers as an instrument instead serves as a direct proxy for our continuous treatment of interest: the probability of eligibility for coverage when single. As such, although the key measure is similar, because the treatment of interest is actually different, both our estimands of interest and our identification assumptions differ.

Let $d_i = Pr(Eligible_i) = F(w_i^s \leq m)$, where w_i^s describes a stochastic income process for individual i if they were single (superscript s), $F(y_i)$ 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. Without specifying the individual income process, we can observe that this value will depend on the features of the income process. 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.

While it provides a starting point for thinking about what we wish to estimate, d_i is not a very interesting object in practice: neither we as the econometricians nor individuals making decisions can perfectly observe it. To make progress, we make the assumption that individual i belongs to a group a of similar individuals, such that $w_i^s \approx w_{a(i)}^s$. Intuitively, we can think about this as expressing that individuals know by observing other, similar individuals roughly what the likelihood is of attaining particular income levels for individuals like themselves.

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. Since individuals in group a face the same income process, within this group $d_{a(i)}$ is the same for any given level of m. In the next section we will see that there is a natural sample analogue for estimating $d_{a(i)}$ based on this assumption.

In a continuous setting, treatment effects are described by the dose-response function:

$$\Delta_i = Y_i(d_i(m^*) - Y_i(d_i(m)) \tag{1}$$

We consider certain restrictions on heterogeneity in treatment effects. We assume that treatment effects can differ for each group a, but not within groups, so $\Delta_{a(i)} = \Delta_i$. Moreover, we assume that the dose-response function for each group a is linear, meaning 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 provides a starting point for further comparison in robustness checks.

Finally, we assume that $m \perp\!\!\!\perp (Y_i|a(i))$. In practice, this 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; and second, that states change policy in response to economic or political conditions that may be correlated with marital outcomes.

The population average treatment effect is the expectation of u_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_{it}, we can write it as follows:

$$\mathbb{E}[\delta_a] = \mathbb{E}[\frac{Y_i(d_i(m^*)|M = m^*, a) - Y_i(d_i(m)|M = m^*, a)}{d(m^*) - d(m)}] \tag{2}$$

The assumption that the probability of eligibility $(d_i(m)|M = m_t, a) = (d_i(m_t)|M = m_t^*, a)$ combined with the independence of Y_i and m imply:

$$\mathbb{E}[Y_i(d_i(m)|M = m^*, a)] = \mathbb{E}[Y_i(d_i(m)|M = m, a)] \tag{3}$$

Since the latter term is observed, it can be substituted into Eq. 2, 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.

3.1.1 Threats to identification

As mentioned above, there are two primary issues with the identifying assumption of exogeneity: endogeneity of migration by individuals and endogeneity of Medicaid expansion by states.

State conditions. The identifying assumption would be violated if, for example, the key measure were correlated with local economic conditions that might also affect marital outcomes. For this reason, although individual earnings are certainly affected by local labor markets, we exclude local variation in earnings as a source of variation in the key measure by defining the earnings distribution at the national level.

In our baseline specifications, we make the assumption that state changes in Medicaid eligibility policy are also uncorrelated with local conditions that may be correlated with the outcomes of interest, as is typical in the application of simulated instruments. However, policy endogeneity is a possible problem. If the timing of changes in eligibility rules is related indirectly or directly to patterns in marital outcomes, then this key assumption would be violated.

Our main specification includes all changes to Medicaid eligibility that occurred between 2011 and 2019. Changes that were made not owing to the ACA Medicaid expansion may plausibly be related to state-specific economic or political conditions. It is straightforward to exclude this source of variation by only using changes owing to decisions to opt into Medicaid expansion: all states that chose never to expand are assigned the same eligibility thresholds that they had in 2010 across all years. However, there is still the additional potential problem that selection into the ACA Medicaid expansion is highly correlated with Democratic control of government institutions. If more Democratic states have systematically different patterns of marital outcomes, identification of the effects of expanded Medicaid eligibility would be confounded.

To test whether these two possible problems matter for results, as a robustness check we employ a strategy based on identification of causal effects with non-random exposure to shocks **?**. The core insight is that, while most Democratic-controlled states expanded, and most Republican-controlled states did not, we can assume exchangeability within each of the groups. In other words, condititional on partisanship, the decision to expand was as-good-as-random. This identifies the effect of increases in Medicaid eligibility only within groups of Democratic- and Republican-controlled states.

Migration. Another potential threat to the identifying assumption would be if individuals moved to (or from, though this seems less plausible) states with more generous Medicaid eligibility levels, and in particular if the characteristics of individuals who migrated in such a way were correlated systematically with marriage patterns. For example, one could plausibly imagine that it is those of higher socioeconomic status that are more mobile, and that they might also have lower marriage and divorce rates than individuals of lower socioeconomic status.

To assess this issue, in robustness checks we test directly whether higher eligibility is associated with net inflows of a given demographic group into a state. Furthermore, we can both directly control for and restrict the sample by whether individuals have recently migrated from another state.

Having described the key points for identification and possible issues therein, the next section discusses data.

3.1.2 Data and variable construction

We use data from the American Community Survey between 2010 and 2019, accessed via IPUMS (Ruggles et al., 2018). We include adults between the ages of 18 and 64 in our sample, thus excluding those who would otherwise be covered by Medicare.[5] The ACS includes information on marital status, as well as information on transition in marital status from the previous year for select marital statuses (including transition into marriage and divorce). The ACS also includes 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 they reside.[6]

To determine eligibility rules by state and year, we use information gathered from yearly reports between 2010 and 2019 by the Kaiser Commission on Medicaid and the Uninsured (see Appendix B for details). We code eligibility levels in terms of the state Medicaid programs or programs that provided equivalent levels of coverage (waiver programs as a result of the ACA also allowed states to establish programs with lesser coverage, which are excluded from this analysis). Eligibility levels for Medicaid vary within states by parental status, income as a percent of the federal poverty line (which in turn depends on family size and income), and work status. However, we use only the eligibility threshold levels for working individuals, as work status is endogenous and the ACA expansion is intended to apply for a given level of income regardless of work status.[7]

To calculate the key variable of interest, the probability that an individual would become eligible for Medicaid in case (s)he is not married, we randomly draw a fixed national 10% sample from the ACS in 2010, after restricting the sample to adults between 18 and 64. Next, we drop individuals who are married. We partition the sample into groups by sex, 5-year age-groups, and 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, excluding local variation in earnings.

[5]Those below the ages of 26 also have the option of coverage via a parent's private insurance policy as a result of Obamacare, 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

[6]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 that depends on family income and family size.

[7]In practice, certain states have received waivers to implement variations on the federal legislation including for work requirements.

Table 1: Summary Statistics for Full Sample

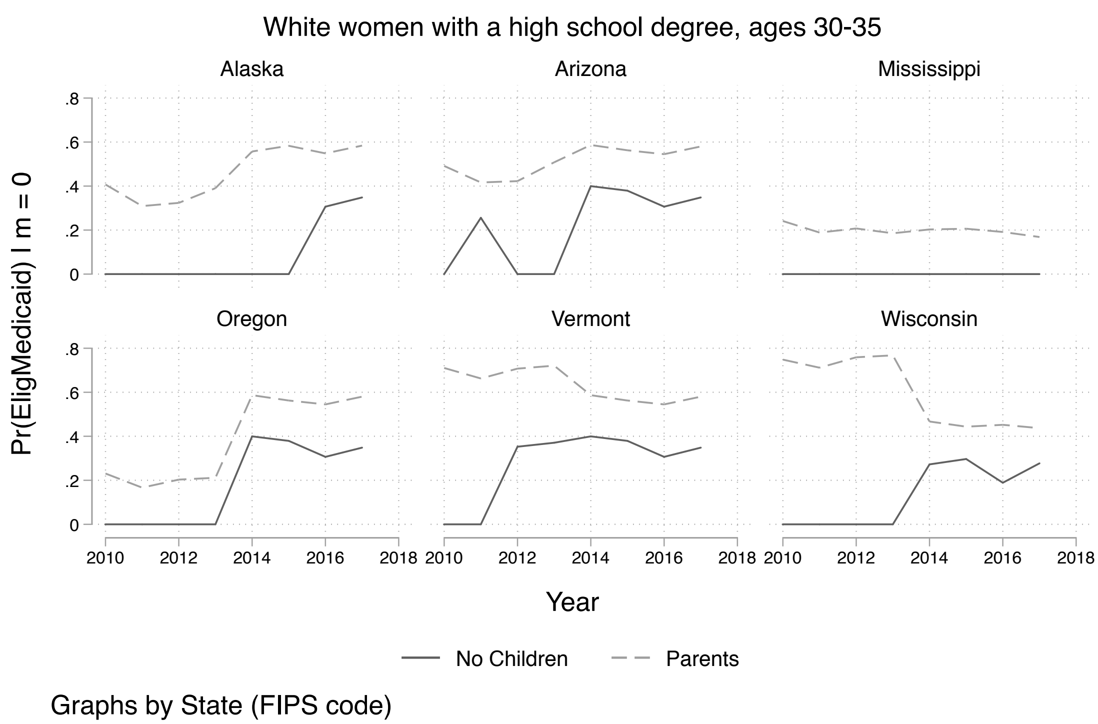
	Men		Women	
	Mean	SD	Mean	SD
Age	41.57	(13.81)	42.19	(13.77)
More than high school	0.51	(0.50)	0.59	(0.49)
Parent	0.36	(0.48)	0.45	(0.50)
Medicaid Coverage	0.07	(0.09)	0.09	(0.10)
Pr(Medicaid)_t)	0.12	(0.13)	0.17	(0.18)
Married	0.49	(0.50)	0.52	(0.50)
Divorced	0.10	(0.30)	0.12	(0.33)
Marriage Rate	0.04	(0.20)	0.04	(0.20)
Divorce Rate	0.02	(0.13)	0.02	(0.13)
Observations	8505124		8706215	

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 weighted by person-specific survey weights. This fraction, $PrMedicaid_{d,s,t}$ (where d indicates the demographic group) is the instrument we use to proxy for the unobserved probability of Medicaid coverage when not married.

Table 1 provides descriptive statistics on the full sample of adults. They have a mean age of 41.6 and 42.2 for men and women respectively. Women are nearly 10 percentage points more likely to be classified as parents, which reflects the fact that mothers most often have primary custody in the case of parental separation. 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 about 2 per 100 couples.

To give some intuition for the identifying variation exploited, Fig. 2 graphs $Pr\left(Medicaid\right)_{d,s,t}$ for two demographic groups, showing the estimated eligibility for six states for both parents and childless individuals who are female, between the ages of 30 and 35, and with a high school degree. The states chosen highlight the fact that state policy variation cannot always be easily grouped into "expansion" and non-expansion states. Among the states pictured, Oregon and Mississippi are the only states that fit into simple definitions of expansion and non-expansion: Mississippi has not changed its eligibility rules at all over the time period studied, and Oregon changed its rules once in 2014 for both parents and non-parents. In comparison, though Alaska is a late-expansion state, officially expanding by September 2015 (and as such, it's coded as expanding for non-parents in 2016) Alaska expanded eligibility for parents in 2014. Arizona and Vermont were both early-expansion states, but Arizona temporarily froze enrollment before completing expansion in 2014, while Vermont in 2014 increased eligibility thresholds for non-parents but slightly decreased it for parents.

Figure 2: Medicaid eligibility for high-school educated women between 30-35



The figure depicts $Pr\left(Medicaid\right)_t$ for a particular demographic group across an array of states in order to illustrate the identifying variation used. $Pr\left(Medicaid\right)_t$ is constructed by calculating the fraction of individuals with the corresponding characteristics in a fixed national sample that would be eligible for Medicaid under a state's rules in a given year.

Finally, Wisconsin is a state that has ostensibly never participated in Medicaid expansion, yet in 2014 both increased eligibility thresholds for its Medicaid-comparable coverage for non-parents and decreased it for parents, such that it has nearly as high eligibility thresholds as mandated by the ACA.

3.2 Estimation

In this section, we describe how we aim to estimate the estimands of interest, given our identification assumptions.

3.2.1 Empirically testing for the effect on new marriages

Intuitively, expanded insurance affects individuals' outside options. How does this affect decisions prior to and at the time of marriage? In the first part of our analysis, we consider whether the Medicaid expansion affects marriage rates as well as whether it affects divorce rates for new marriages, depending on the generosity of Medicaid at the time of marriage.

Empirical Prediction 1: Decreased marriage rate. As non-married individuals' outside options improve, they become pickier about potential spouses. As a result, we expect to observe relatively fewer transitions from being single to married where Medicaid eligibility is higher.

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

$$\text{MarrLastYear}_{i,s,t} = \eta Pr\left(Medicaid\right)_{d,s,t} + \gamma_d + \mathbf{X}_{i,s,t}\beta + \varepsilon_{i,s,t} \tag{4}$$

The equation is estimated for women and men separately. The subscript d indicates a demographic group, s the state in which they reside, and t the year. The dependent variable is the marriage rate for the demographic group in a given state and year. Given the data we use, this is calculated by taking the fraction of individuals who report being married within the last year relative to the population of individuals that were not married last year, that is to say the sum of those who were married in the last year and those who report being currently non-married with no marital status change in the last year.

The key variable of interest is $Pr\left(Medicaid\right)_{d,s,t}$, 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 (e.g. each group defined by age, parental status, and education) and $\mathbf{X}_{i,s,t}$ represents a vector of additional control variables. The vector of control variables includes full sets of indicator variables for year and state, as well as the fully set of indicator for demographic groups. Additional specifications include 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, as well as state-level controls variables for the party affiliation of the governor for local economic conditions and measures of political control. Standard errors are clustered at the state level. We now turn to a second empirical prediction which corresponds to our first theoretical prediction.

Empirical Prediction 2: Decreased incidence of divorce for higher eligibility at time of marriage. If individuals become pickier about who they will marry, a natural consequence is that the couples that do still get married will be more positively selected. Thus, we would predict that divorce rates for marriages formed under greater Medicaid generosity should be lower.

To test this prediction, we instead assign $Pr\left(Medicaid\right)_{d,s,t^*}$ and associated control variables for t^*, where t^* indicates the year of marriage. Given that our data spans 2010 to 2017, we restrict the sample to those individuals married in this time span. While the year of marriage is observed, we need to make certain inferences about other aspects of eligibility to assign the correct level. Age at the time of marriage, and hence the correct 5-year age group, can of course be easily calculated. If an individual is observed to have a child older than the length of the current marriage, we infer that the individual was also parent at time t^*. If the individual's oldest child is currently younger than the length of the marriage, we instead infer that the individual was not a parent at the time of marriage. This latter inference may be mistaken given that individuals only report the age of the oldest child in the household. If the individual was indeed a parent at prior to the time of marriage, but the oldest child has since moved out of the household, we would underestimate eligibility at the time of marriage, since eligibility levels for parents are in all cases equal to or (more commonly) more generous than for childless individuals.

We make the simplifying assumptions that there is no change in educational attainment since the time of marriage and that individuals have not moved states. With respect to education, since we consider a relatively short time period and most of the sample is past typical schooling ages, this assumption is not particularly restrictive. If anything, errors in this assumption would have also have the effect of underestimating eligibility at the time of marriage, since individuals can only previously have had lower educational attainment (and hence a lower earnings expectation).

Assuming individuals have not moved is potentially problematic. A particular concern may be that people of higher socioeconomic status, who might have lower divorce rates, might be more mobile and hence more likely to move to high-eligibility states. During the time period of our sample, we have information only on whether an individual has moved in the past year, so we cannot entirely exclude people who have may have moved during the observed time period. However, it is possible to examine whether there is systematic migration towards higher-eligibility states, and this approach and associated results are discussed later in the paper.

We associate $Pr\left(Medicaid\right)_{d,s,t^*}$ given these inferred characteristics at the time of marriage. The eligibility levels are otherwise calculated exactly as before.

We estimate the following equation in relation to this prediction:

$$\text{DivLastYear}_{i,s,t} = \eta Pr\left(Medicaid\right)_{d,s,t^*} + \gamma_d + \mathbf{X}_{i,s,t^*}\beta + \alpha_{t-t^*} + \varepsilon_{i,s,t} \tag{5}$$

The structure of the equation is similar to Eq. 4 although the right-hand side variables are with respect to characteristics at the time of marriage, as well as indicator variables for the years elapsed since marriage to account for common trends in the likelihood of divorce over the course of a marriage.[8]

3.2.2 Empirically testing for the effect on intra-household transfers

A final natural question is to explore how couples that are already married are affected in terms of divorce rates by Medicaid expansion.

To answer this question, we estimate the following equation:

$$\text{DivLastYear}_{i,s,t} = \eta Pr\left(Medicaid\right)_{d,s,t} + \gamma_d + \mathbf{X}_{i,s,t}\beta + \varepsilon_{i,s,t} \tag{6}$$

This equation bears similarities to both Eq. 4 and Eq. 5. The right-hand side is identical to Eq. 4, though the outcome is with respect to divorce rates. While it does consider an effect of Medicaid eligibility on divorce rates, the distinction from Eq. 5 is that we estimate the *contemporaneous effect* of a change in eligibility on couples that were already married when a change went into effect.

These empirical predictions are written in their most general form, and therefore do not directly mention potential (interesting) heterogeneity across demographic characteristics. However, these might obviously play an important role. To address these important (potentially) heterogeneous effects, we will, for each of the analyses, also consider heterogeneity by age and educational attainment level, which are included by interacting $Pr(Medicaid)_{d,s,t}$ with the relevant categorical variable.

4 Results

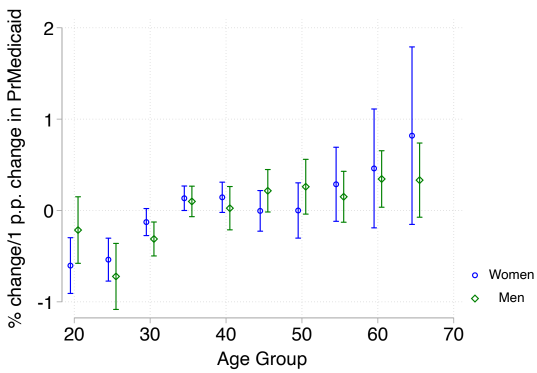
We now turn to a discussion of the empirical fidings. In summary, we find strong support for our empirical predictions (and thereby indirectly for our theoretical predictions), together with both some more obvious and also more surprising heterogeneous patterns. In particular, we indeed find that increased eligibility for Medicaid decreases marriage rates for individuals in their 20s and 30s, even among individuals with higher levels of educational attainment. Additionally, it decreases the rate of divorce for individuals married under higher eligibility, but has no evident net effect on divorce rates of already-married individuals. We now discuss each of these results in detail.

4.1 Marital Outcomes

We start with the results of estimating Eq. 4 for men and women, by age groups. In particular, Fig. 3 shows estimates of marginal effects by age group using Eq. 5 as a starting point, where age group is fully interacted with $Pr\left(Medicaid\right)$. The effects in this figure are scaled relative to each group's baseline marriage rate such that each represents the percent change in the marriage rate associated with a 1 p.p. increase in $Pr\left(Medicaid\right)$. This figure immediately highlights the fact that there are negative effects to marriage rates in response

[8] After controlling for indicators for the year of marriage and the years since marriage, we can no longer include calendar year dummy variables, similarly to the standard age-cohort-period issue of collinearity.

Figure 3: The effect of increased eligibility on marriage rates, by age



Including individuals between the ages of 18 and 64 from the 2010 to 2019 waves of the ACS, this graph presents 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 age group to yield age-group specific marginal effects. Each regression includes fixed effects for demographic group, year, state, and time-varying demographic group economic conditions as well as the state governor's party affiliation.

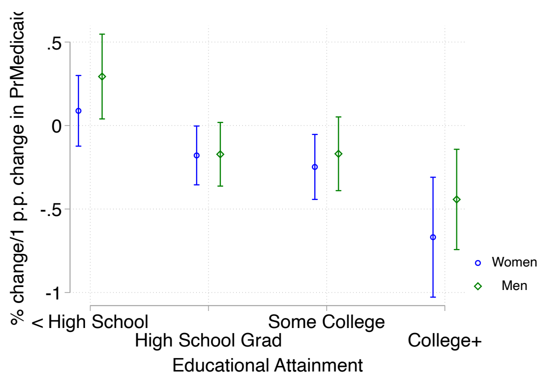
to increases in the likelihood of Medicaid eligibility, specifically for younger individuals in their 20s and 30s. In contrast, there seems to be almost no effect. This is quite intuitive. Most individuals make their first marriage decisions when they are younger adults and one plausible explanation could be that individuals postpone entry into marriage, which is also compatible with the theoretical model.

Beyond age, differences across education levels might also be very relevant in assessing the effects of increases in the likelihood to become eligible for Medicaid. In particular, education levels could be correlated with expected earnings, particular marriage markets they face (i.e., their feasible set of potential spouses) and possibly even risk aversion. Furthermore, the value of marriage might also be affected by the education levels of the partners, e.g. by the dependency of the surplus function on the educational composition of the couple. To provide some insight on how these types of differences might impact marriage responses, we perform a similar estimation as for age but with educational attainment levels, where education is considered a proxy for all these other factors. Fig. 4 displays these results, which like the previous figure rescales each subgroup by dividing the coefficient by the group's baseline marriage rate. The results suggest that individuals who have dropped out of high school have either no significant response (women) or actually a positive response (men). Those with high school degrees and some college have small decreases of about a quarter of a percent for a 1 p.p. increase in $Pr\,(Medicaid)$ which are similar across men and women, although only significant at the 10% level. Intriguingly, college educated individuals have the largest negative responses, with point estimates around or above 0.5%.

This last result highlights an important new fact: even individuals with high educational attainment, who we think of normally as having high expected earnings, may exhibit behavioral responses to the implementation or expansion of means-tested programs. This finding is consistent with our theoretical discussion that these programs might affect behavior either through individuals directly gaining coverage or indirectly through a perception of an increased likelihood to be covered, in the event of a shock. Since many studies of means-tested public programs focus on recipients and other signifiers of high likelihood of enrollment like low educational attainment, this result suggests it may be useful to consider effects of such programs on a broader basis.

To be clear, the absolute increases in the likelihood of eligibility for those with a college degree are lower than for other groups, so the net effect on marriages is in fact smaller, yet in proportional terms the response is larger.

Figure 4: The effect of increased eligibility on marriage rates, by educational attainment



Including individuals between the ages of 18 and 64 from the 2010 to 2019 waves of the ACS, this graph presents coefficients from a regression of marriage in the past 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 age group to yield age-group specific marginal effects. Each regression includes fixed effects for demographic group, year, state, and time-varying demographic group economic conditions as well as the state governor's party affiliation.

Finally, the average estimated effect of increased Medicaid eligibility for the whole population does not appear to be significantly different from zero, as illustrated in Table 2, which presents results from estimating Eq. 4 for men and women. For both men and women, the estimated effect of an increase in $Pr\,(Medicaid)$ on the marriage rate is small and not significantly different from zero. Indeed, for men four decimal points are needed to observe a non-zero digit. This further adds to the point that the effects of increasing the likelihood to become eligible for Medicaid is concentrated in younger adults who are the most likely to be active on the marriage market.

The corollary of the prediction on lower expected marriage rates was that subsequent divorce rates of newly married couples might be lower under more generous Medicaid regimes (cfr. Empirical Prediction 2.) Table 3 shows the average effect on divorce rates for the whole poulation, examining whether higher eligibility levels at the time of marriage affect subsequent divorce rates. Though it has a similar structure to Table 2, it's important to keep in mind that the right-hand variables are defined (or inferred) with respect to the

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

Marr. Rate in t	Men			Women		
$Pr(Medicaid)_dt$	-0.000	0.000	0.000	-0.003	-0.003	-0.003
	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)
$Unemp.Rate_{dt}$		-0.046	-0.045		-0.021	-0.020
		(0.008)	(0.008)		(0.009)	(0.009)
$LogWage_{dt}$		0.009	0.009		0.010	0.010
		(0.002)	(0.002)		(0.002)	(0.002)
$NotinLF_{dt}$		-0.033	-0.033		-0.001	-0.002
		(0.004)	(0.005)		(0.006)	(0.006)
$DemGovernor_{st}$			0.000			0.000
			(0.000)			(0.000)
Observations	4,473,268	4,473,177	4,458,707	4,365,340	4,364,990	4,348,270
Effect Size in Percent	0	.001	.001	-.007	-.006	-.006
Local Market Controls			YES			

Results are estimated by OLS using adults ages 18 to 64 from the American Community Survey, waves 2010-2019. 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 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.

year of marriage. In the baseline specification in Cols. 1 and 3, the estimated coefficient on $Pr\,(Medicaid)_t *$ is close to zero and not statistically different from zero. However, once controls for contemporaneous state conditions at the time of the interview are included, both coefficients become more negative and statistically significant. In particular, these controls include the mean unemployment rate and log wage for each demographic group. This latter specification implies a 10 p.p. increase in Medicaid eligibility at the time of marriage leads to a 3.4% reduction in divorce rates over the first nine years of marriage for men, and a 1.6% reduction for women.

The difference in these specifications implies that conditional on controlling for other economic factors that may push couples towards divorce and that vary over time for demographic groups, greater insurance eligibility at the time of marriage reduces the subsequent risk of divorce.

Additionally, given that we can examine the effect of eligibility at the time of marriage only on a short time span, the cumulative effect over the full course of these marriages may be larger. To provide some hint as to whether this is the case, Fig. 5 displays the coefficient on $PrMedicaid$ for the specification including controls for state conditions, separately by

Table 3: The effect of the probability of Medicaid eligibility at the time of marriage on the rate of divorce

Divorced in $t*+i$	Men			Women		
$Pr(Medicaid)_{dt^*}$	-0.001	-0.006	-0.006	-0.002	-0.003	-0.003
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$Unemp.Rate_{dt}$		0.094	0.095		-0.000	-0.003
		(0.007)	(0.008)		(0.008)	(0.008)
$LogWage_{dt}$		0.001	0.001		-0.003	-0.003
		(0.001)	(0.001)		(0.001)	(0.001)
$NotinLF_{dt}$		0.112	0.113		-0.067	-0.068
		(0.003)	(0.003)		(0.003)	(0.003)
$DemGovernor_{st}$			-0.001			0.001
			(0.001)			(0.001)
Observations	883,058	883,036	880,436	888,164	888,105	885,730
Effect Size in Percent	-.05	-.34	-.34	-.096	-.16	-.17

Standard errors in parentheses

Results are estimated by weighted least squares, where the weights are given by the number of individuals in each demographic group, using adults ages 18 to 64 who were married between 2010 and 2017, from the American Community Survey, waves 2010-2019. Whether an individual was divorced in the prior year is regressed on the probability of Medicaid eligibility at the time of marriage given the individual's demographic characteristics and year of marriage.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, all defined relative to the time of marriage), 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.

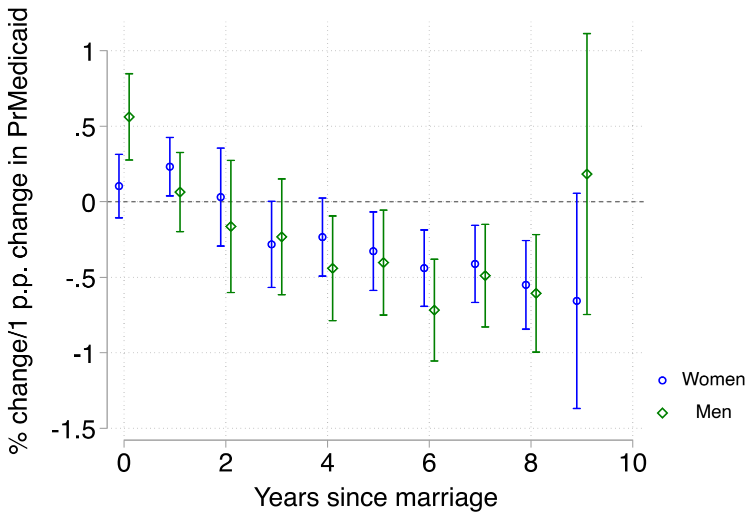
years since marriage for men and women. There is a clear pattern that the coefficients become more negative with years of marriage, and this pattern is particularly pronounced for men.

Together, these results on marriage and divorce provide indirect evidence that individuals in marginal relationships postpone or avoid marriage, which results in higher quality–and hence lower divorce rates–of the remaining marriages that still do occur.

4.2 Divorce rates for already-married individuals

We have also posited that while eligibility at the time of marriage may affect divorce rates of newly formed couples, changes to eligibility may also affect the rate of divorce for already-formed couples. Table 4 presents results pertaining to this hypothesis, again displaying specifications including the baseline controls (Cols. 1 and 4) as well as with additional controls in remaining columns. For already-married men, there appears to be a small and statistically significant effect of Medicaid eligibility on divorce, although in terms of effect

Figure 5: The effect of the probability of Medicaid eligibility at the time of marriage on the rate of divorce, by years since marriage



Results are estimated by OLS using adults ages 18 to 64 from the American Community Survey, waves 2010-2019. Whether an individual was divorced in the prior year is regressed on the probability of Medicaid eligibility at the time of marriage given the individual's demographic characteristics and year of marriage. The probability of eligibility is interacted with indicators for years since marriage to estimate the marginal effect for each year since marriage. 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, all defined relative to the time of marriage), 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. Coefficients are divided by each group's baseline divorce rate to yield the percent change in divorce rate for a 1 p.p. increase in $PrMedicaid_t*$.

size the impact is small: it implies a 10 percentage point increase in Medicaid eligibility would decrease divorce rates by only 0.14

Unlike for marriage rates, considering different dimensions of possible heterogeneity does not reveal any significant effects for subgroups. Figs. 6 and 7 display the results by age and education, again scaled by baseline rates of divorce for each subgroup. The point estimates for individuals in their late 20s and early 30s are negative and smaller than for any other group, which is suggestive, but there are no statistically significant effects. Likewise for education, the point estimates are centered close to zero.

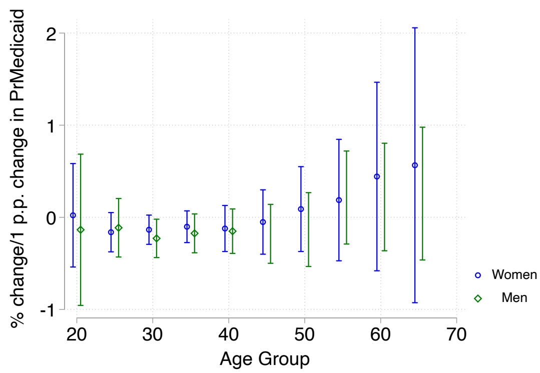
The lack of a positive effect on contemporaneous divorce perhaps goes against intuition. However, this result itself highlights the importance of considering general equilibrium effects, since the prediction that expanding Medicaid should increase divorce is based on partial equilibrium thinking, where other factors are kept constant. One important note about these null effects is that they are not necessarily an indication of *no effect* of expanded Medicaid eligibility, since our theory points to the possibility for its effect to either decrease or increase divorce rates, depending on individual conditions for a couple. These overall null

Table 4: The effect of the probability of Medicaid eligibility on contemporaneous divorce rates

Divorced in $t-1$	Men			Women		
$Pr(Medicaid)_{dt}$	-0.002	-0.002	-0.002	-0.001	-0.001	-0.001
	(0.002)	(0.001)	(0.001)	(0.002)	(0.001)	(0.001)
$Unemp.Rate_{dt}$		0.038	0.040		0.050	0.049
		(0.007)	(0.007)		(0.007)	(0.007)
$LogWage_{dt}$		-0.002	-0.002		0.001	0.001
		(0.001)	(0.001)		(0.002)	(0.002)
$NotinLF_{dt}$		0.026	0.026		-0.007	-0.007
		(0.005)	(0.005)		(0.003)	(0.003)
$DemGovernor_{st}$			-0.000			0.001
			(0.000)			(0.000)
Observations	4,468,064	4,468,001	4,461,657	4,760,778	4,760,570	4,754,109
Effect Size in Percent	-.014	-.014	-.015	-.008	-.007	-.008

Results are estimated by OLS using adults ages 18 to 64 from the American Community Survey, waves 2010-2019. 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.

Figure 6: The effect of increased eligibility on divorce rates, by age



Including individuals between the ages of 18 and 64 from the 2010 to 2019 waves of the ACS, this graph presents coefficients from a regression of divorce rates 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 age group to yield age-group specific marginal effects. Each regression controls for demographic group fixed effects, year, state, and local state economic and political conditions. Coefficients are divided by each group's baseline divorce rate to yield the percent change in divorce rate for a 1 p.p. increase in $PrMedicaid_t*$.

effects may thus mask true treatment effects that may go in either direction for individual couples.

4.3 Robustness

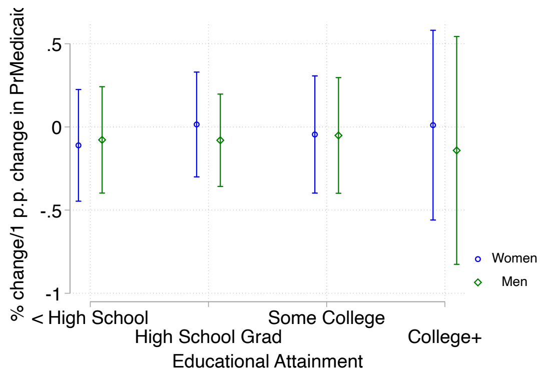
As outlined in the description of the empirical strategy, there are a few potential concerns with assuming that the Medicaid policy variation exploited is exogenous. In this section we examine the robustness of the core results to the potential issues of endogenous migration and a correlation between selection into Medicaid expansion and local economic or political conditions.

4.3.1 Migration

With respect to migration, the primary concern would be that individuals choose to migrate to states with more generous Medicaid policies. Such a movement would be particularly problematic in case such individuals happen to have lower marriage and divorce rates, since then their characteristics and endogenous migration might be driving our observed results, which we interpret as indicative of the effects of the policy change itself. Results are presented in Appendix C.

The ACS data for the time period we study includes information on whether individuals have moved from another state in the past year. As a starting point, we first use this variable as a dependent variable in regression of the same form as used throughout the paper. Specifically, we consider whether an individual moved states in the past year as the

Figure 7: The effect of increased eligibility on divorce rates, by educational attainment



Including individuals between the ages of 18 and 64 from the 2010 to 2019 waves of the ACS, this graph presents coefficients from a regression of divorce rates 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 educational attainment to yield marginal effects by education level. Each regression controls for demographic group fixed effects, year, state, and local state economic and political conditions. Coefficients are divided by each group's baseline divorce rate to yield the percent change in divorce rate for a 1 p.p. increase in $PrMedicaid_t*$.

dependent variable of a regression equation that is otherwise identical to the right-hand sides in Eq. 4 and Eq. 6. We carry out this estimation for the full sample as well as restricted to individuals under the age of 26 (to address the fact that the largest effects on marriage rates come from this younger group). Interestingly, for both the full and younger samples, $Pr\,(Medicaid)$ is positively associated with having moved to the current state of residence from a different state last year for men, while for women there is no apparent relationship. This suggests that migration may be a cofounder.

To address whether this is the case, we test two strategies. First, we add a control for the share that moved in a given demographic group, state and year. Second, we exclude all individuals that report moving from the sample. We replicate our age-specific and education-specific estimates of $Pr\,(Medicaid)$. The only notable difference in the results is that when recent movers are excluded, marginal effects for the age groups spanning approximately age 50-64 become more positive and statistically significant at the 5% level. If anything then, this would suggest that recent movers are more likely to divorce.

4.3.2 Policy endogeneity

Another potential concern is that at least some state-level expansions of Medicaid eligibility may actually be in response to changes in economic or political conditions, which might in turn be related systematically to marital outcomes. One specific example for the latter concern is that, if Democratic-leaning states have a higher prevalence of more culturally liberal attitudes (e.g., with respect to marriage and in particular divorce), and these states are also more likely to expand Medicaid more considerably, then we are actually capturing a spurious correlation between $Pr\left(Medicaid\right)_t$ and marriage outcomes.

To address this concern, we replace $Pr\left(Medicaid\right)$ with a reconstructed variable where demographic groups of a given state are initially assigned values based on 2010 levels of Medicaid eligibility. This level of eligibility is only replaced for changes owing to a state's decision to opt into Medicaid expansion, thus excluding local changes that are unconnected to the national Medicaid expansion. Next, states are split by whether or not their governor was a Democrat in 2011. Within this group, the key assumption is that the decision to expand is effectively as good as random. To account for differing trends, the regression includes a full set of interaction variables between the Democratic status of the governor in 2011 and year fixed effects.[9] Results for these additional analyses are presented in Appendix D.

The restructuring of the key variable does not appear to meaningfully affect the coefficients on $Pr(Medicaid)$. Similarly to the robustness checks for migration, this restricted version of the key variable likewise does not meaningfully affect the results–the youngest groups still exhibit declines in marriage and a clear gradient by education, as well as lowered divorce rates over the course of marriage given higher levels of eligibility at the time of marriage. Interestingly, the only difference is again that the oldest three groups in fact exhibit increases in contemporaneous divorce.

5 Theoretical Framework

5.1 The environment

In this section, we present a stylised model that highlights how Medicaid expansion might affect the marriage market, and allows for rationalizing results of the previous section. The model is most closely related to the analysis presented in Anderberg (2007), who studies intra-household risk sharing contracts within a search-based equilibrium of the marriage market. Though we will focus on Medicaid as the specific empirical setting in this paper, the theoretical results are more broadly generalizable to other public insurance programs, shielding individuals against all sorts of negative income shocks.

The fundamentals of our model are as follows: there is a continuum of men (m) and women (f). We assume an equal number of men and women, each of which has a measure normalized to one. Time is continuous and individuals discount the future by a common discount factor r.

In each moment of time, every individual faces uncertainty regarding the (indirect) utility (s)he will receive, due to the possibility to receive an adverse shock. Whether or not an individual receives a shock is indicated by ω, where $\omega = -1$ in case (s)he does face a financial shock and $\omega = 1$ in case (s)he does not. We assume that the utility flow for each individual depends on (i) marital status and (ii) the realization of the shock. The following gives an overview of the utility flows across both:

[9] In the spirit of **?**, this interaction also picks up the expected rate of expansion in a given year for a given group.

Table 5: Utility flows across marital status and ω.

	Married	Single
$\omega = -1$	$\psi(-1)$	$\psi_0(-1)$
$\omega = 1$	$\psi(1)$	$\psi_0(1)$

Hence, individual i receives utility $\psi^i(\omega)$, in case i is married and the corresponding utilities in case (s)he is single is denoted by $\psi_0(\omega)$. [10] The shock $\omega = -1$ indicates an exogenous loss in (indirect) utility. Given the stylized nature of our model, this shock can have multiple interpretations. The simplest way one can think of the realization $\omega = -1$ is that the individual faces an exogenous reduction in resources, e.g. a negative income shock.[11]. The situation where $\omega = 1$ denotes the baseline case. The associated probability of receiving a shock $\omega = -1$ is denoted by π.

Match quality and intra-household transfers We now discuss the two main motivations for marriage in our model: the match quality (non-material benefit) and transfers which provide insurance. Whenever two single individuals meet they draw a match quality θ from a (known) distribution G. Match quality can change over time. In the baseline version of the model, we assume that the match quality process is *memoryless*, that is, in each period of time there is a probability λ that the match quality changes, and a new match quality is drawn from the same distribution G.[12]

Utilities of spouses are assumed to be perfectly transferable. After match quality has been revealed, the spouses agree on (state-contingent) transfers to redistribute utilities within the household,[13] in particular, at each moment in time the husband agrees to pay the wife a transfer equal to $t(\omega;\theta)$. The latter depends on the particular environment the couple is facing, that is the vector of shocks $\omega = \left(\omega^m, \omega^f\right)$ and the match quality, θ. The expected utility flow for the husband is therefore equal to:

$$\tilde{\psi}^m - \tilde{t}$$

and for the wife:

$$\tilde{\psi}^f + \tilde{t},$$

[10] Notice that we are making the assumption here that all individuals of the same gender have the same preferences.

[11] Given that the focus of our paper is on the effects of Medicaid expansion as a public insurance scheme, another way to think about $\omega = -1$ is that the individual is more exposed to unexpected high medical expenses, which lowers his/her indirect utility. Notice that we do not explicitly model consumption, savings, or other intra-household allocations such as labor supply, we only require here that $\omega = -1$ lowers the utility flow for an individual.

[12] The memoryless assumption is useful to simplify the analysis and is not uncommon in search-equilibrium models of the marriage market, e.g. Shin (2015) and Goussé et al. (2017). We will discuss robustness of our main results to more persistent match quality processes later on in this section.

[13] Hence, these transfers constitute a source of intra-household insurance.

where $\tilde{\psi}^i = \pi\psi^i(-1) + (1-\pi)\psi^i(1)$, and $\tilde{t}$ denotes the expected transfer from the husband to the wife. The (expected) transfers are determined in order to split the marital surplus between the spouses (cfr. infra).

Search on the marriage market. The random process of match quality imposes a risk of divorce on each spouse. In particular, whenever the match quality becomes too low, the relative benefits of staying inside the marriage decreases and each spouse might unilaterally decide to break up the match. After returning to the pool of singles, they can start searching for a new spouse. We will use S to denote the measure of singles at any point in time. We assume that finding a new partner is costly in the sense that it might take time to find a new partner.[14] To be more precise, two opposite-sex singles can meet each other with a probability[15] $\phi(S)$, where $\phi(0) = 0$ and $\phi'(S) > 0$ for all $S \in (0,1)$.

5.2 Equilibrium

We now solve for the equilibrium on the marriage market. First, the continuation value for a single individual i is given by

$$r\mathcal{V}_0^i = \tilde{\psi}_0^i + \phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{V}^i\left(\tilde{\theta}\right) - \mathcal{V}_0^i, 0\right\}\right]. \tag{7}$$

The first term in (7) represents the (expected) utility flow of an individual at a given moment in time, in case i meets a potential partner (with probability $\phi(S)$), there are two possibilities: either the value of matching with the proposed partner (and receiving match quality $\tilde{\theta}$) is larger than the value of remaining single, $\mathcal{V}_0^i$, in which case (s)he will accept the match. Otherwise, (s)he rejects the proposed match and remains single. The continuation value for a man i who is currently married and the associated match quality is equal to θ is given by:

$$\begin{aligned} r\mathcal{V}^m(\theta) &= \tilde{\psi}^m - \tilde{t} + \tfrac{1}{2}\theta \\ &+\lambda\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{V}^m\left(\tilde{\theta}\right) - \mathcal{V}^m(\theta), \mathcal{V}_0^m - \mathcal{V}^m(\theta)\right\}\right]. \end{aligned} \tag{8}$$

At each moment in time a husband receives the (expected) utility flow, pays expected transfers $\tilde{t}$ and enjoys the match quality of the current marriage. With probability λ he faces a shock in the match quality of his marriage, after which he has to compare the value of staying with his current wife at the new match quality $\tilde{\theta}$ with his outside value of divorcing and becoming single. Furthermore, in order to obtain the expected continuation value, the husband also has to take expectations over the possible realization over the utility shock ω. A similar expression can be obtained for the value of a married woman:

[14] Given the strong symmetry present in our model, e.g., equal gender ratio and the homogeneity of preferences within each gender, a frictionless equilibrium model of the marriage market would be less appropriate to explain the co-existence of singles and married individuals with positive match quality.

[15] We follow Anderberg (2007) and assume, for simplicity, that singles can only meet one potential spouse.

$$
\begin{aligned}
r\mathcal{V}^f(\theta) &= \tilde{\psi}^f + \tilde{t} + \tfrac{1}{2}\theta \\
&+\lambda\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{V}^f\left(\tilde{\theta}\right) - \mathcal{V}^f(\theta), \mathcal{V}_0^f - \mathcal{V}^f(\theta)\right\}\right]. \qquad (9)
\end{aligned}
$$

Given that we have assumed utilities are perfectly transferable within marriage, we can easily obtain an expression for the marital surplus of a marriage with current match quality θ, $\mathcal{S}(\theta)$:

$$\mathcal{S}(\theta) = \mathcal{V}^m(\theta) + \mathcal{V}^f(\theta) - \mathcal{V}_0^m - \mathcal{V}_0^f.$$

This marital surplus is then split according to (generalized) Nash bargaining. At any moment in time the expected transfer solves:

$$\tilde{t} = \arg\max\left(\mathcal{V}^f(\theta) - \mathcal{V}_0^f\right)^{\gamma}\left(\mathcal{V}^m(\theta) - \mathcal{V}_0^m\right)^{1-\gamma}.$$

This implies the following:

$$\mathcal{V}^m(\theta) = \mathcal{V}_0^m + (1-\gamma)\,\mathcal{S}(\theta), \qquad (10)$$

for married men and

$$\mathcal{V}^f(\theta) = \mathcal{V}_0^f + \gamma\mathcal{S}(\theta), \qquad (11)$$

for married women. Hence, each individual gets their outside value (value of singlehood) and a share of the marriage surplus, where the share is given by their relative bargaining weight. Using (7), (8)-(9) and (10)-(11), we can write down the continuation value for the marriage surplus (for a match with match quality level θ):

$$
\begin{aligned}
(r+\lambda)\,\mathcal{S}(\theta) &= \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f + \theta \\
&+\lambda\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right] \\
&-(1-\gamma)\,\phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right] \\
&-\gamma\phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right]. \qquad (12)
\end{aligned}
$$

We note that $\mathcal{S}(\theta)$ is increasing in θ, hence, we can easily characterize the acceptance strategy of singles. In particular, we need to find the lowest possible match quality such that an individual is indifferent between accepting the potential spouse and remaining single and search for a spouse. Such a cut-off match quality, which we will denote by $\underline{\theta}$ can be found by solving:

$$\mathcal{S}(\underline{\theta}) = 0,$$

which can be simplified to the following expression[16]:

$$\underline{\theta} + \frac{\lambda}{r+\lambda}\varphi\left(\underline{\theta}\right) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f = \frac{\phi}{r+\lambda}\varphi\left(\underline{\theta}\right). \tag{13}$$

The first step to solve for an equilibrium is to find a unique solution $\underline{\theta}$ for (13), conditional on the pool of singles, S. Our first result is that this is always possible[17]:

Proposition 5.1 (cut-off match quality)**.** *Given G, λ, π, r, S and ϕ, there exists a unique $\underline{\theta}$ which satisfies:*

$$\underline{\theta} + \frac{\lambda - \phi\left(S\right)}{r+\lambda}\varphi\left(\underline{\theta}\right) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f = 0. \tag{14}$$

And each individual will accept a spouse if $\theta \geq \underline{\theta}$.

Given a solution $\underline{\theta}$ to (14), we can find the corresponding size for the pool of singles. Indeed, we will follow Anderberg (2007) and Goussé et al. (2017) and focus on the steady state of our model, in which case the flow outside of marriage and into marriage should be equal. Reformulating this, we obtain the following expression for flow equilibrium:

$$\phi\left(S\right)S\left[1 - G\left(\underline{\theta}\right)\right] = \lambda\left(1 - S\right)G\left(\underline{\theta}\right). \tag{15}$$

The left hand side represents the flow into marriage, which is equal to the fraction of those singles meeting each other ($\phi\left(S\right)S$) and drawing a match quality which is higher than the cut-off level $\underline{\theta}$. The right hand side is the flow out of marriage, which is given by the fraction of married individuals who draw a new match quality below the cut-off level. Equation (15) indeed states that in steady state and for a given level $\underline{\theta}$ these flows should be equal. Furthermore, (15) implicitly defines a function $\underline{S} = \underline{S}\left(\underline{\theta}\right)$.[18] We will impose that, in steady state equilibrium, the pool of singles we obtain given a value for $\underline{\theta}$ should be then consistent with the latter cut-off match quality, this can be more precisely formulated:

Definition 5.2 (Marriage market equilibrium)**.** *Given the primitives of the model, $\left\{\psi^i\left(\omega\right), \psi_0^i\left(\omega\right), \pi, G, \lambda, \varphi, r\right\}$, a marriage market equilibrium consists in reservation match quality, $\underline{\theta}$ and measure of singles, $\underline{S}$, such that:*

1. *Given $\underline{S}$, the value of $\underline{\theta}$ is a solution to (14) and*

2. *The value of $\underline{\theta}$ produces a solution, $\underline{S}$, for (15) that yields a match probability consistent with $\underline{\theta}$.*

[16] Notice that $\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right] = \mathbb{E}_\pi \int_{\underline{\theta}}^{\infty} \mathcal{S}\left(\tilde{\theta}\right) dG\left(\tilde{\theta}\right) = \mathbb{E}_\pi \frac{1}{r+\lambda}\int_{\underline{\theta}}^{\infty}\left[1 - G\left(\tilde{\theta}\right)\right] d\tilde{\theta}$. Where we used partial integration for the second equality. Now, if we define $\varphi\left(\underline{\theta}\right) = \mathbb{E}_\pi \int_{\underline{\theta}}^{\infty}\left[1 - G\left(\tilde{\theta}\right)\right] d\tilde{\theta}$, and substitute this into the equation $\mathcal{S}\left(\underline{\theta}\right) = 0$, we obtain (13).

[17] We refer to the Appendix for a proof of this and all subsequent results.

[18] It can be shown that (15) defines a unique solution, $\underline{S}$ for a given $\underline{\theta}$.

Hence, more succinctly stated, in order to find a marriage market equilibrium, we need to find a fixed point of the composite mapping

$$\underline{S}\left(\underline{\theta}\left(\phi\left(.\right)\right)\right).$$

Notice that ϕ is increasing in S, from (14) it can be shown that $\underline{\theta}$ is increasing in ϕ and from (15) we can show that $\underline{S}$ is increasing in $\underline{\theta}$. Existence of a marriage market equilibrium is then a direct result from Tarski's fixed point theorem. Summarizing:

Proposition 5.3 (existence of a marriage market equilibrium)**.** *Given the primitives of the model, $\{\psi^i(\omega), \psi_0^i(\omega), \pi, G, \lambda, \varphi, r\}$, a marriage market equilibrium $(\underline{\theta}, \underline{S})$ always exists.*

An important remark here is that, even though Proposition 5.3 shows the existence of a marriage market equilibrium, it does not imply uniqueness of an equilibrium, as there generally will be multiple equilibria[19]. However, if we further assume that the support for match quality is sufficiently large, then there exist $-\infty < \theta^L < \theta^U < +\infty$ such that everyone will reject the proposed match with match quality $\theta < \theta^L$ and conversely, everyone would like to stay in a match with $\theta > \theta^U$. This guarantees that all the equilibria are interior in the sense that $S \in (0, 1)$.

5.3 The role of public insurance

We now turn our attention to the role and effects of public insurance on the marriage market. Suppose that a social planner (government) introduces a transfer scheme[20] to help insure individuals against adverse shocks, i.e., the contingency in which $\omega = -1$. The transfer amount is denoted by $\tau > 0$. The eligibility to such a transfer is assumed to depend on marital status. To be more precise, when an individual i is single (s)he receives a transfer in case $\omega = -1$, hence (s)he receives as expected transfer $\pi\tau$. In contrast, married individuals receive in expectation transfers equal to $\pi \times \pi \times \tau$, that is, married individuals are only eligible to receive a public transfer in case *both* spouses receive a negative shock, i.e., $\omega^m = \omega^f = -1$.[21] The first step to study how public insurance affects the marriage market equilibrium is to study how the acceptance strategies of individuals are impacted by the introduction of the public transfer scheme. Adjusting the indifference condition between accepting a potential spouse and staying single, we obtain:

[19] This is a general feature for search-based equilibria models of the marriage market and it also applies to the frameworks presented in Anderberg (2007) and Goussé et al. (2017).

[20] The way we model a public insurance scheme is not too dissimilar to the approach taken by French and Jones (2011) and Capatina (2015) to model Medicaid. In particular, they assume Medicaid operates as a consumption floor, i.e., a minimum level of consumption which the government guarantees to all individuals.

[21] The fact that we restrict the amount of expected public transfers to couples is consistent with many means-tested social programs, and in particular such restrictions are also present in the Medicaid program, where eligibility is based in part on the family income and household structure. Eligibility of individuals depends fundamentally on the characteristics of the household overall and not only his/her own circumstances. In the context of our model we assume that individuals only obtain access to public transfers depending on the realization of ω for both spouses.

$$\underline{\theta} + \frac{\lambda - \phi}{r + \lambda} \varphi\left(\underline{\theta}\right) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f + \pi\left(\pi - 2\right)\tau = 0. \tag{16}$$

The difference between (14) and (16) is the presence of the term $\pi(\pi - 2)\tau$, which is negative. The condition in (16) implicitly defines a function $\underline{\theta}(\tau)$. We can then easily show the following result[22]:

Lemma 5.4 (New cut-off match quality). *Assume $\lambda \leq \phi(S)$, then $\underline{\theta}(\tau)$ is non-decreasing.*

In order to obtain a definitive monotonicity result, we impose the (sufficient, not necessary) condition that $\lambda \leq \phi$. This condition holds in case either the likelihood to find a new partner (ϕ) is sufficiently large or match quality is sufficiently stable (i.e. the likelihood to receive a shock to match quality is small). The intuition of such an assumption is that, while it's realistic to imagine that couples do experience shocks to the quality of their relationship from time to time, such shocks in already formed relationships should not be more frequent than a single individual meeting a potential match (regardless of quality).

As a consequence, individuals will become pickier in terms of accepting new partners in the face of higher outside options. The result in Lemma 5.4 is very intuitive: the public transfer scheme offers a substitute to one of the main benefits of being in a partnership: insuring each other against shocks. In order to still enter or stay in marriage, the non-material benefit (i.e., match quality) has to increase.

Turning to the overall impact of the public transfers on the marriage market, we need to study the effect of τ on the equilibrium measure of singles. One complication is the potential presence of multiple equilibria. Given this, we will focus on the behaviour of the lowest and highest equilibrium measure of singles[23],denoted by $\underline{S}^L(\tau)$ and $\underline{S}^U(\tau)$ respectively. To make progress, we first observe that the result in Lemma 5.4, together with the monotonicity (and continuity) of $\underline{S}\left(\underline{\theta}\left(\phi\left(.\right)\right)\right)$, implies that the latter (composite) mapping is also increasing in τ. We can then immediately apply Corollary 1 from Milgrom and Roberts (1994) and obtain:

Proposition 5.5 (Monotonicity of equilibrium measure of singles). *Suppose that $\lambda \leq \phi$, then $\underline{S}^L(\tau)$ and $\underline{S}^U(\tau)$ both increase in τ.*

Hence, we can conclude that the introduction of the public transfer scheme shifts the set of equilibrium measures of singles 'upwards'.

5.4 Extensions

Though the framework presented so far is relatively stylised and omits several aspects which might be of importance, it already provides some specific results in terms of the (expected) effects of public insurance on the marriage market. We now discuss several extensions of our basic model, in order to add more realism and derive further implications.

[22] See the Appendix for a derivation of this result.

[23] In the literature on monotone comparative statics, it is customary to refer to these equilibria as 'extremal equilibria', e.g., Topkis (1978), Milgrom and Roberts (1994) and Amir (2005).

5.4.1 Persistent match quality process.

The model presented so far assumed that match quality follows a memoryless process. In reality, it might be more plausible to assume that match quality follows a more persistent process in which the likelihood to enjoy a good match quality in the future is higher the better the match quality is today. As an example, Anderberg (2007) studies a version of his model in which match quality is a discrete random variable following a Markov process and in which he imposes a first-order stochastic dominance-type condition on the transition matrix, that is, the cumulative distribution for tomorrow's match quality is an increasing function in today's match quality level.[24] Notice that, as a consequence of the memoryless assumption on match quality, when the critical match quality, $\underline{\theta}$ increases, the divorce rate of newly formed couples also increases since the latter is given by $\lambda G(\underline{\theta})$.

We would predict that, as a consequence of the public transfer scheme, the divorce rate for newly wed couples would increase due to the fact that individuals are pickier and given that shocks are purely random, the likelihood to fall below this increased match quality-standard is higher. In contrast, if we would adapt our model to allow for a persistent match quality process, the divorce rate for newly married couples would be predicted to be lower given the fact that it's now more likely to receive a higher match quality over time, which is due to the fact that individuals are pickier, and therefore select themselves in relatively higher match quality matches, which then remain of 'high quality' for some amount of time due to the persistence in match quality.

5.4.2 Direct effects of public transfers.

So far we have assumed that the expected indirect utilities $\tilde{\psi}^i$ and $\tilde{\psi}_0^i$ are not directly affected by the introduction of public insurance. Clearly, this is overly restrictive, especially in a richer model where we would explicitly model these indirect utilities as the result of labor supply and consumption choices. The mere presence of public insurance might affect these choices and hence have an additional effect on the indirect utilities, that is

$$\frac{\partial \tilde{\psi}^i}{\partial \tau} \neq 0,$$

and

$$\frac{\partial \tilde{\psi}_0^i}{\partial \tau} \neq 0.$$

Given the means-tested nature of many public insurance program at the household level, it is intuitive to assume that the direct effects on utility are larger for singles than married individuals, in particular, we can assume that:

$$\frac{\partial \tilde{\psi}_0^i}{\partial \tau} - \frac{\partial \tilde{\psi}^i}{\partial \tau} \geq 0, \text{ for } i = m, f.$$

By inspection of (16), it is then easily seen that the result in Lemma 5.4 remains true.

[24] Another popular alternative to generate persistence in match quality is to assume it follows a random walk, see e.g., Mazzocco et al. (2014) and Low et al. (2018).

5.4.3 Intra-household allocations.

The couple was assumed to split the marital surplus through generalised Nash bargaining, as highlighted by equations (10) and (11). We can back out the (expected) transfers paid by the husband to the wife, which gives us the following:

$$\tilde{\psi}^f + \frac{1}{2}\theta + \tilde{t} = r\mathcal{V}_0^f + \gamma\left(\tilde{\psi}^m + \tilde{\psi}^f + \theta - r\mathcal{V}_0^f - r\mathcal{V}_0^m\right),$$

hence,

$$\tilde{t} = (1-\gamma)\, r\mathcal{V}_0^f - \gamma r\mathcal{V}_0^m + \gamma\tilde{\psi}^m - (1-\gamma)\,\tilde{\psi}^f + \left(\gamma - \frac{1}{2}\right)\theta. \tag{17}$$

The expected transfer, $\tilde{t}$ are increasing in the outside value of women and decreasing in the outside value for men. In the baseline version of our model, the introduction of public insurance affects husbands and wives symmetrically, and there is no change in the expected intra-household transfers. Given our favourite interpretation of ω as a negative income shock, which then leads an individual being more exposed to high medical expenses, it might be more realistic to assume that public transfers to men and women are different, that is, $\tau^f \neq \tau^m$. Such an alternative assumption can be rationalised by observing that on average women face higher medical expenditures than men (see e.g. Cylus et al. (2010)).[25] For ease of exposition, let $\tau^f = \tau$ and $\tau^m = \kappa\tau$ with $\kappa \in [0, 1]$ a fixed parameter. We can then show the following result:

Proposition 5.6 (Expected intra-household transfers)**.** *The expected transfer from m to f increases as long as*

$$\left(\frac{\partial\tilde{\psi}_0^f}{\partial\tau} - \frac{\partial\tilde{\psi}^f}{\partial\tau}\right) \geq \kappa\frac{\gamma}{1-\gamma}\left(\frac{\partial\tilde{\psi}_0^m}{\partial\tau} - \frac{\partial\tilde{\psi}^m}{\partial\tau}\right). \tag{18}$$

Note that (18) requires some direct effect of public insurance on the indirect utility flows. The condition is more easily satisfied if κ is small, which is consistent with the value of transfers being much larger for women than men. Furthermore, public insurance is more likely to increase intra-household insurance from husband to wife in case her initial share of the total surplus is relatively smaller. This is intuitive given that there are not many possibilities to further increase expected transfer amounts in case the wife already has a very large share of surplus. Finally, expected transfers from husband to wife will increase in case the relative increase in the wife's indirect utility is larger, which is more intuitive given that $\tau^f > \tau^m$.

We should further add that, in the absence of explicit intra-household decisions (e.g., labor supply, consumption, household production or childcare decisions), this is the only

[25] An alternative assumption might be to allow for different probabilities to face an adverse shock, i.e., $\pi^m \neq \pi^f$. This could be due to differences in employer-sponsored health insurance, for example because married women are more likely to be out of the labor force or work part-time. However, such differences in likelihood to face such shocks opens up the channel of labor force participation and job decisions, which is beyond the scope of the present paper.

effect of the reform. If we would include such richer intra-household choice aspects, the predicted effect of a change in transfers from one spouse to the other can also have further effects. For example, if we assume that leisure is a normal good, then the increased (expected) transfers from husband to wife would imply an increase in wives' leisure or simultaneously a decrease in labor supply. Hence, if the public insurance scheme affects individuals in a different manner (here: differences across genders), the public transfers associated with the insurance will have broader effects in terms of allocations within the household.

5.4.4 Savings and public goods.

In the model as we have presented so far, we have assumed that individuals' only incentive to marry each other has the purpose of risk sharing and to obtain idiosyncratic non-material benefits (match quality), which we have refered to as 'love'. In that way, we have abstracted from several other features which might be of importance to study matching in the marriage market. For example, we omitted included storage of resources in the form of wealth/savings in the model, which clearly matters in terms of a household's ability to cope with financial shocks (e.g. uninsured medical expenses).[26] However, including wealth/savings seriously complicates the analysis of the search-based equilibrium of the marriage market, since in that case the acceptance strategies for individuals (i.e., which potential partners are deemed acceptable for marriage) would then not only depend on the match quality, but also on the amount of assets both potential spouses have at their disposal. More importantly, we also neglected another main reason to enter marriage, namely to produce and enjoy public goods, most notably children's welfare. When this motive is taken into account, it naturally opens up the possibility that the acceptance strategies of individuals will depend on their (and their prospective spouse's) education levels. A particular mechanism for this dependence is that parents' education are direct inputs in the production of children's 'quality', e.g. Del Boca et al. (2014) and Chiappori et al. (2017). An important implication of such models is that there will be positive assortative matching on education.[27] We could incorporate the production of public goods within a match, by explicitly allowing for marital surplus to depend on the education levels of both spouses. This channel, we would suggest, would not add much further insight to the question how expansions in public insurance affects the marriage market, as this directly relates to the risk-sharing motive. However, we do want to remark that improving public insurance, ceteris paribus, might have an impact on sorting patterns of individuals along the education dimension. Indeed, in case - in the context of the present model - risk

[26]More generally, even on the macro-economic level there might be interesting links between saving rates and health expenditures, e.g. Chen et al. (2019) have highlighted the relationship between rising health expenditures and declining saving rates in the US context.

[27]Many papers in the empirical matching literature explicitly allow for the marital surplus to depend on the education levels of the spouses, e.g. in the seminal contribution by Chiappori et al. (2006), also see the good overview by Chiappori and Salanié (2016). Another strand of the literature is interested in how (changes in) the assortativeness on education in the marriage market has contributed to income inequality, e.g. Eika et al. (2019) and Chiappori et al. (2020).

(i.e., exposure to the idiosyncratic financial shocks) is correlated with education levels, then the provision of public insurance affects the relative costs and benefits to consider potential matches with a different education level. More specifically, if we return to the interpretation of the shocks in our model as uninsured medical bills pertaining to health shocks, then marrying a lower educated spouse is more 'risky' from a pure risk sharing motive.[28] Hence, an expansion in public insurance therefore reduces, ceteris paribus, the perceived 'riskiness' of lower educated individuals. Furthermore, if assortative mating for higher educated individuals is less pronounced, then improving public insurance (e.g. due to the Obamacare reform) which is not aimed at individuals with higher education might still impact the latter through the increased attractiveness of lower educated individuals as potential spouses.

5.5 Discussion

It is useful to summarize the results we derived from the model to assess how the key predictions relate to the empirical results. In particular we derived the following predictions:

> **Theoretical Prediction 1: increase of match quality.** As a consequence of the introduction of the public insurance scheme, the cut-offmatch quality, $\underline{\theta}$ increases, cfr. Lemma 5.4.
>
> **Theoretical Prediction 2: Increase in expected transfers with ambiguous effects on divorce rates among already-married couples.** In case the introduction of the public transfer scheme affects males and females in a differential way, we have shown that intra-household allocations will be affected, in particular, the spouse whose outside option increases more following the introduction of public insurance will see a compensation through higher transfers from his/her spouse, cfr. Proposition 5.6.
>
> **Theoretical Prediction 3: behavioral responses to prospective eligibility.** The model predicts that both individuals who directly gain insurance (because they are in the unlucky state and single) and those who indirectly gain the possibility of insurance (because they are in the lucky state or are married) will change behavior with the introduction of public insurance.

The first prediction clearly relates to the findings that marriage rates decreased among young adults, while divorce rates were lower for new marriages. The model highlights the intuitive result that these changes are arguably a result of individuals being pickier about the (non-pecuniary) quality of a match, and choosing to forego marriages otherwise. This highlights an important result, particularly in the context of policy discussions targeted at "strengthening" marriage as an institution: there may be tradeoffs between the quality or

[28] There is a large literature listing evidence that higher education (socio-economic status) is strongly positively correlated with good health (hence less exposure to relatively high medical bills), e.g. Adams et al. (2003), Adda et al. (2003), Galama and Van Kippersluis (2019), Yabroff et al. (2019).

stability of marriages and the number of marriages. It should not be taken as a given that it is good for marriage as an institution to incentivize more of them, since it also likely will mean more will end in divorce.

The second prediction helps to rationalize the limited effects we find for changes to divorce of already-married people. While it might seem intuitively that divorce rates should increase if outside options improve, the model highlights that there are many more factors at play and possible net outcomes when we take into account intra-household transfers and the fact that each member of a couple might benefit in different ways from insurance expansion. As a result, it's in fact not strange that we find no net change to overall divorce rates, or at most small increases for older age groups under certain specifications.

Finally, the model provides insights through the third prediction on how the possibility of insurance may affect behavior even in moments when individuals are not currently eligible. This channel helps to provide intuition for the fact that we observe substantial responses for individuals with high educational attainment. Even though they have low absolute probabilities of becoming eligible, they are not negligible. While we demonstrate this for the specific setting of Medicaid, the model highlights that this is arguably a more general mechanism applicable to other types of means-tested public insurance. In turn, it suggests that it may be worth studying how other aspects of public insurance affect outcomes beyond those who specifically gain eligibility or coverage at a particular time of observation.

5.6 Discussion

6 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 Obamacare. In particular, we have show that an increase in the probability of Medicaid eligibility leads to a decrease in the marriage rate among young adults and equally in a decrease in the divorce rate of newly married couples, for those with higher probabilities of eligibility at the time of marriage. These results are consistent with our stylized search-based equilibrium model of the marriage market where match quality increases in response to the introduction of expansion of public insurance that benefits singles. We also provide evidence that there is no net change in divorce rates of already-married couples in response to increased Medicaid eligibility.

Beyond demonstrating the substantial impacts of Medicaid on marital outcomes, an important contribution of our work is to show that expansions of public insurance programs mainly targeted at lower-income individuals may still affect the behavior of those who may have low absolute likelihoods of ever gaining coverage.

There are several avenues for future research. First, we think it would be extremely useful to address more normative aspects pertaining to the overall welfare effects of the Medicaid expansions. Indeed, the present paper has shown that the Obamacare reform had quite large effects on the marriage market, both in terms of potentially reducing entry into marriage, but equally changing the quality of newly formed marriages. The latter clearly has the potential to generate long-run positive welfare effects (e.g. investments in human capital formation of children within newly formed marriages etc.) Another aspect pertains to optimal public insurance design. 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, e.g. Choi and Valladares-Esteban (2020), a similar exercise for our setting could be an interesting new application. Finally, given data restrictions, we have mainly focussed on the effects of the Medicaid expansions on marriage and divorce rates and duration of marriages. However, given our suggestive results regarding broader effects (including individuals with higher education), it would be interesting to study potential compositional changes in the marriage market, e.g. the degree of sorting along education or wealth levels.

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A Proofs

Proof of Proposition 5.1

Starting from (7), and substituting (10) and (11) and $\varphi\left(\underline{\theta}\right) = (r+\lambda)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right]$, we obtain the following:

$$r\mathcal{V}_0^f = \tilde{\psi}_0^f + \frac{\gamma\phi\left(S\right)}{r+\lambda}\varphi\left(\theta\right), \tag{19}$$

for single women and

$$r\mathcal{V}_0^m = \tilde{\psi}_0^m + \frac{\left(1-\gamma\right)\phi\left(S\right)}{r+\lambda}\varphi\left(\theta\right), \tag{20}$$

for single men. Substituting the value functions for singles in (13) yields the following equivalent expression:

$$(r+\lambda)\,\mathcal{S}\left(\underline{\theta}\right) = \tilde{\psi}^m + \tilde{\psi}^f + \underline{\theta} + \frac{\lambda}{r+\lambda}\varphi\left(\underline{\theta}\right) - r\mathcal{V}_0^m - r\mathcal{V}_0^f = 0, \tag{21}$$

which implicitly defines the reservation match quality, $\underline{\theta}$ as a function of $\left(\mathcal{V}_0^m, \mathcal{V}_0^f\right)$. In particular,

$$\underline{\theta}\left(\mathcal{V}_0^m, \mathcal{V}_0^f\right) = \mathcal{F}\left(r\mathcal{V}_0^m + r\mathcal{V}_0^f - \tilde{\psi}^m - \tilde{\psi}^f\right). \tag{22}$$

Substituting $\underline{\theta}\left(\mathcal{V}_0^m, \mathcal{V}_0^f\right)$ in (19)-(20), we obtain the following system:

$$\begin{cases} \mathcal{V}_0^f = \frac{\tilde{\psi}_0^f}{r} + \frac{1}{r}\frac{\gamma\phi(S)}{r+\lambda}\varphi\left(\underline{\theta}\left(\mathcal{V}_0^m, \mathcal{V}_0^f\right)\right) = 0, \\ \mathcal{V}_0^m = \frac{\tilde{\psi}_0^m}{r} + \frac{1}{r}\frac{(1-\gamma)\phi(S)}{r+\lambda}\varphi\left(\underline{\theta}\left(\mathcal{V}_0^m, \mathcal{V}_0^f\right)\right) = 0. \end{cases}$$

Which we can write down succinctly as $\mathcal{V} = \mathcal{G}\left(\mathcal{V}\right)$, where $\mathcal{V} = \left(\mathcal{V}_0^m, \mathcal{V}_0^f\right)$. We now proceed by showing that the above system of equations has a unique solution, $\left(\mathcal{V}_0^{m*}, \mathcal{V}_0^{f*}\right)$. To do this, note that the Jacobian of $\mathcal{G}$ is given by,

$$\mathcal{J}_{\mathcal{G}} = \begin{pmatrix} -\frac{\gamma\phi(S)(1-G(\theta))}{r+\lambda G(\theta)} & -\frac{\gamma\phi(S)(1-G(\theta))}{r+\lambda G(\theta)} \\ -\frac{(1-\gamma)\phi(S)(1-G(\theta))}{r+\lambda G(\theta)} & -\frac{(1-\gamma)\phi(S)(1-G(\theta))}{r+\lambda G(\theta)} \end{pmatrix}$$

Now, notice that

$$\|\mathcal{J}_{\mathcal{G}}\| < 1,$$

and therefore, we can invoke Banach's fixed point theorem to show that there exists a unique $\left(\mathcal{V}_0^{m*}, \mathcal{V}_0^{f*}\right)$ which solves $\mathcal{V} = \mathcal{G}\left(\mathcal{V}\right)$. Substituting in (22) then gives the (unique) cut-off level for match quality.

Proof of Lemma 5.4: Given that $\pi(\pi - 2) < 0$, we must have that $\underline{\theta} + \frac{\lambda-\phi}{r+\lambda}\varphi(\underline{\theta})$ increases. If $\lambda \leq \phi$, and because $\frac{\partial\varphi(\underline{\theta})}{\partial\underline{\theta}} < 0$, it must be that $\underline{\theta}$ increases.

Proof of Proposition 5.6: From (7), (10) and (11) we obtain the following expression for the value of singlehood:

$$r\mathcal{V}_0^f = \tilde{\psi}_0^f + \phi(S)\gamma\varphi(\underline{\theta}), \tag{23}$$

and

$$r\mathcal{V}_0^m = \tilde{\psi}_0^m + \phi(S)(1-\gamma)\varphi(\underline{\theta}). \tag{24}$$

Using these and substituting in the expression for expected transfers, (17), we obtain:

$$\tilde{t} = (1-\gamma)\left(\tilde{\psi}_0^f - \tilde{\psi}^f\right) - \gamma\left(\tilde{\psi}_0^m - \tilde{\psi}^m\right) + \left(\gamma - \frac{1}{2}\right)\theta \tag{25}$$

Differentiating and noting that $\tau^m = \kappa\tau$ for some fixed $\kappa \in [0,1]$, we get that $\frac{\partial\tilde{t}}{\partial\tau} \geq 0$ if and only if (18) is satisfied.

B 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 Medciaid-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.

C Robustness Checks: Migration

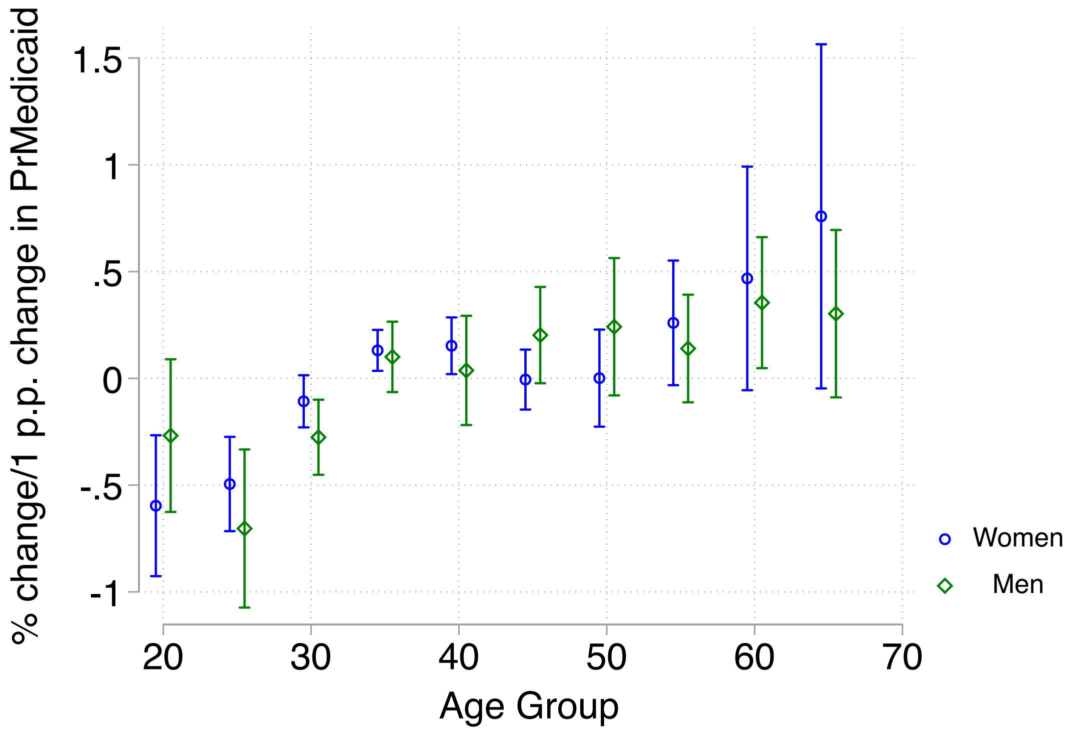
Table C.1: The effect of the probability of Medicaid eligibility on likelihood of moving states

Moved States since $t-1$	Men		Women	
$Pr(Medicaid)_{dt}$	0.004	0.009	-0.001	-0.003
	(0.002)	(0.004)	(0.002)	(0.004)
$Unemp.Rate_{dt}$	0.093	-0.053	0.073	-0.010
	(0.020)	(0.039)	(0.025)	(0.037)
$LogWage_{dt}$	-0.007	0.004	0.001	0.020
	(0.005)	(0.011)	(0.006)	(0.012)
$NotinLF_{dt}$	0.005	0.096	0.028	0.094
	(0.012)	(0.033)	(0.013)	(0.035)
$DemGovernor_{st}$	0.000	0.001	0.000	0.000
	(0.000)	(0.001)	(0.000)	(0.001)
Sample	All	<Age 30	All	<Age 30
Observations	8,485,472	2,122,306	8,684,168	2,039,238

Results are estimated by weighted least squares, where the weights are given by the number of individuals in each demographic group, using adults ages 18 to 64 from the American Community Survey, waves 2010-2017. The dependent variable is the divorce rate for the demographic group, namely the fraction of individuals reporting divorce in the last year relative to the total number of individuals that were married one year earlier. The controls in Cols. 1 and 3 include dummies for year, state, educational attainment (four levels), five-year age groups, race (white, black, and other races combined as a third group) and parental status. Cols. 2 and 4 additionally allowe for interactions between the full set of year dummies and education, parental status, age group, and race. Standard errors are clustered at the level of parental status and state.

Figure C.1: Effects of Eligibility on Marriage by Age

(a) Controlling for the fraction of movers



(b) Excluding recent movers

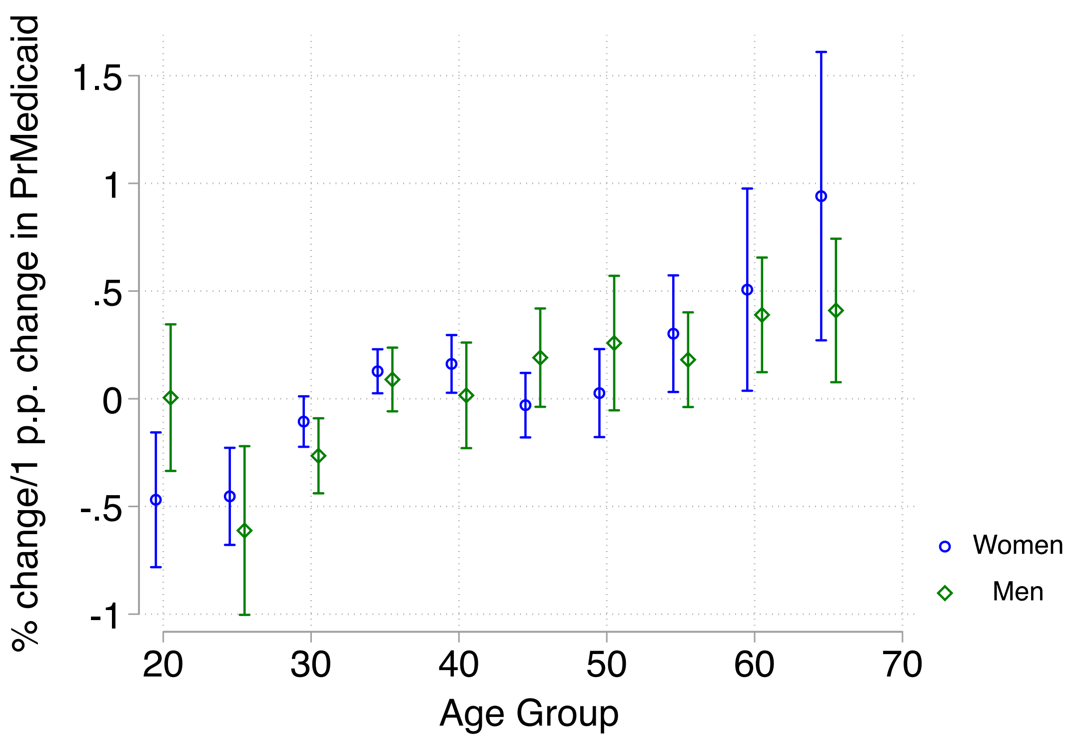
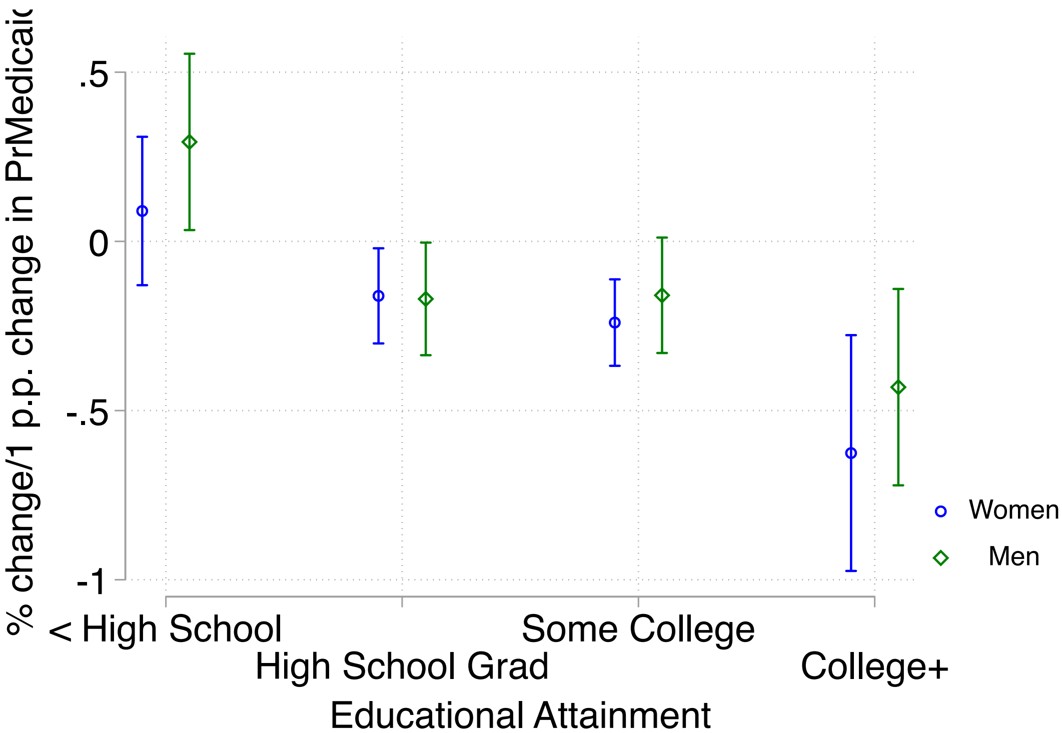


Figure C.2: Effects of Eligibility on Marriage by Education

(a) Controlling for the fraction of movers



(b) Excluding recent movers

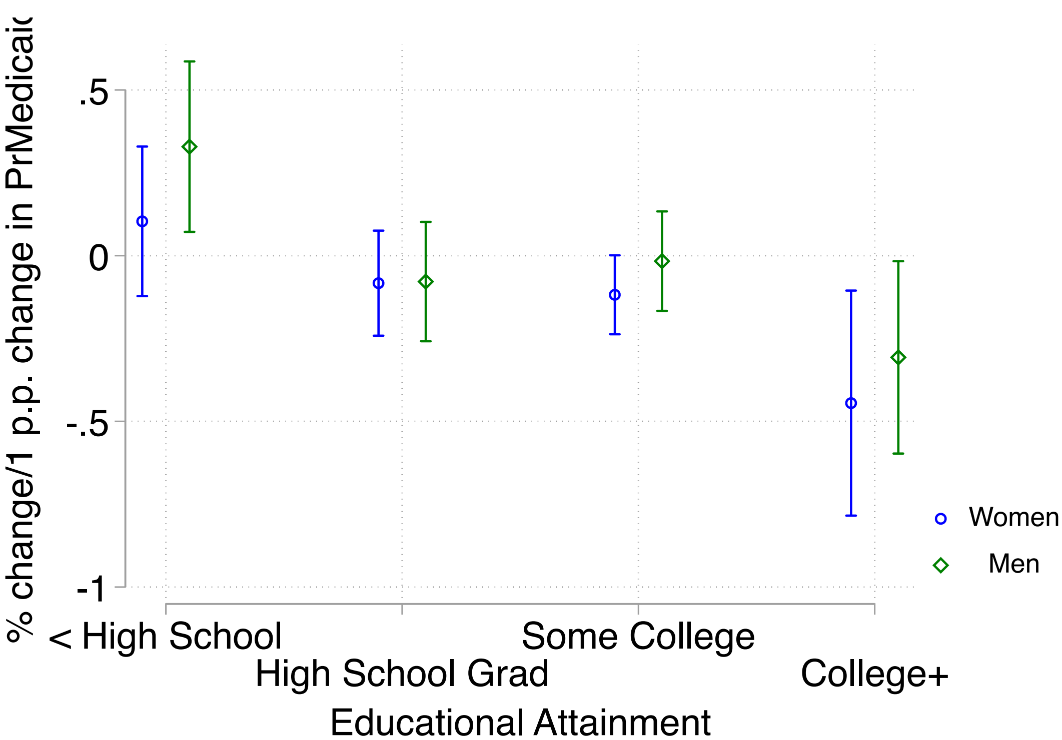
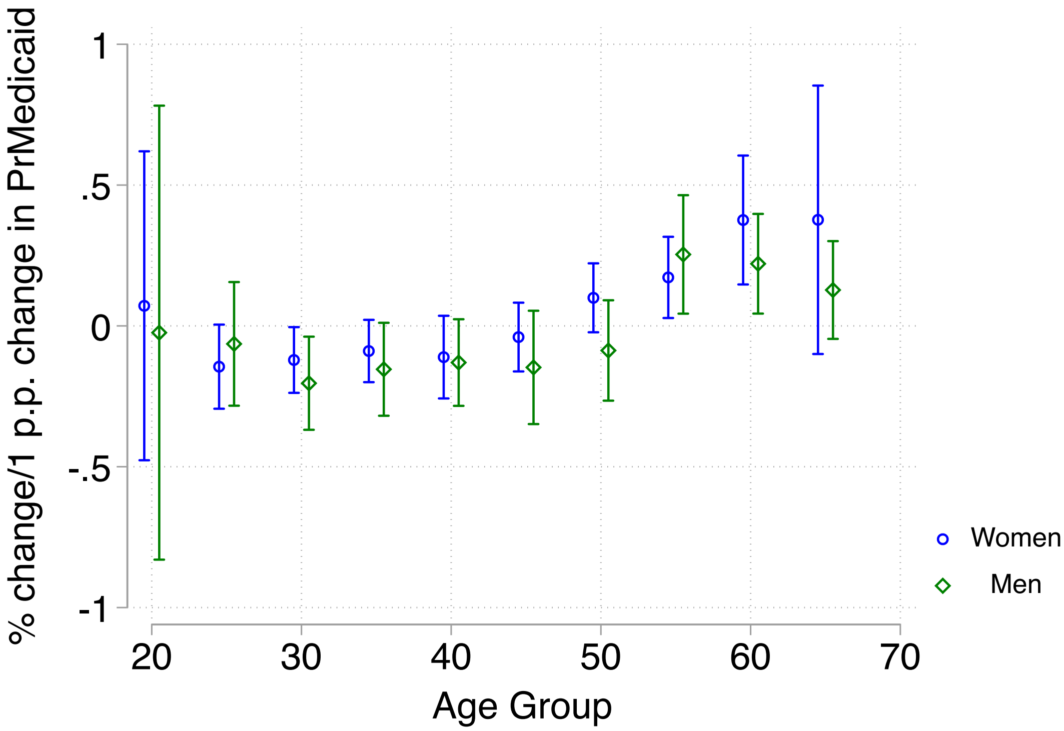


Figure C.3: Effects of Eligibility on Contemporaneous Divorce by Age

(a) Controlling for the fraction of movers



(b) Excluding recent movers

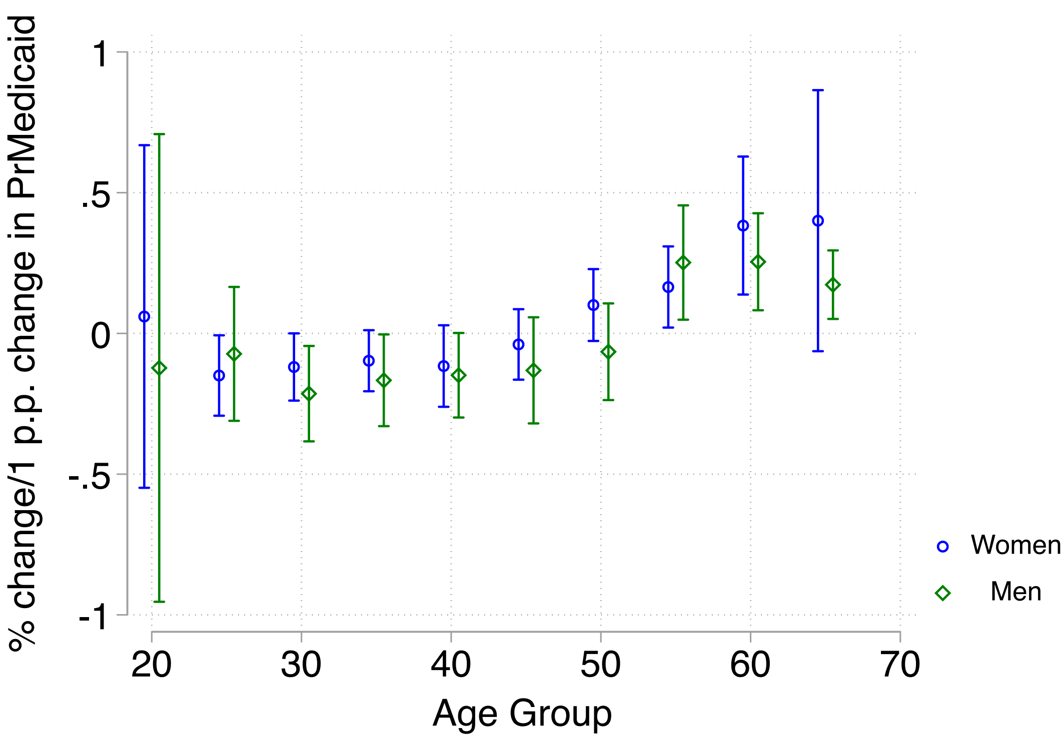
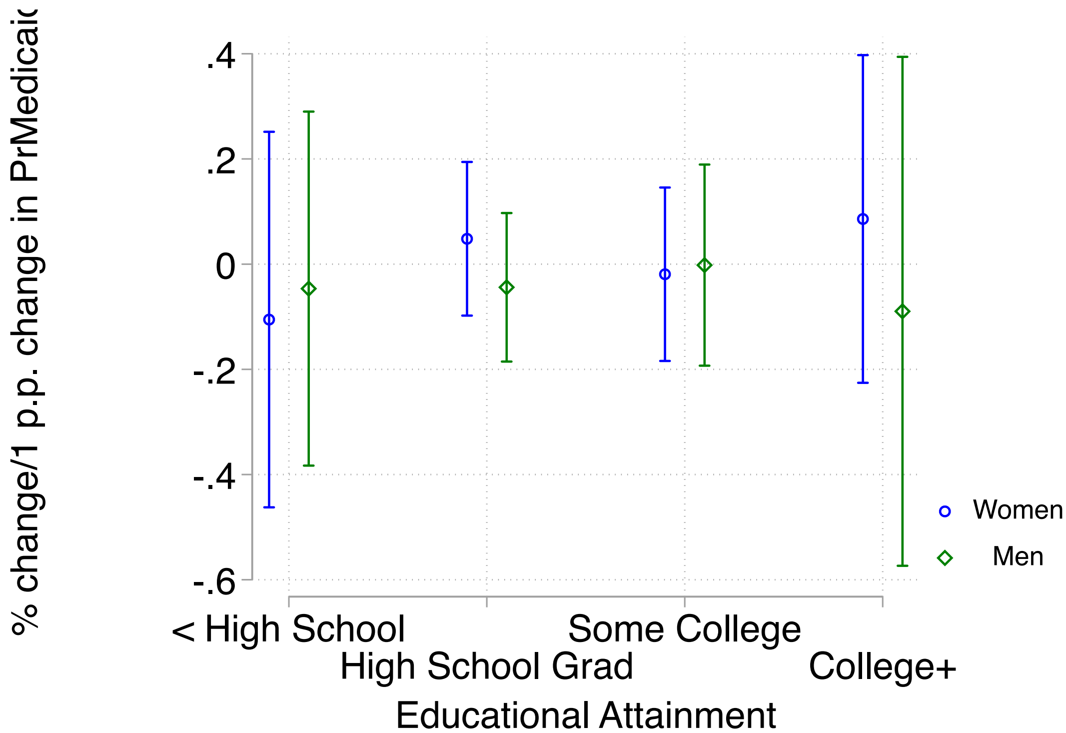
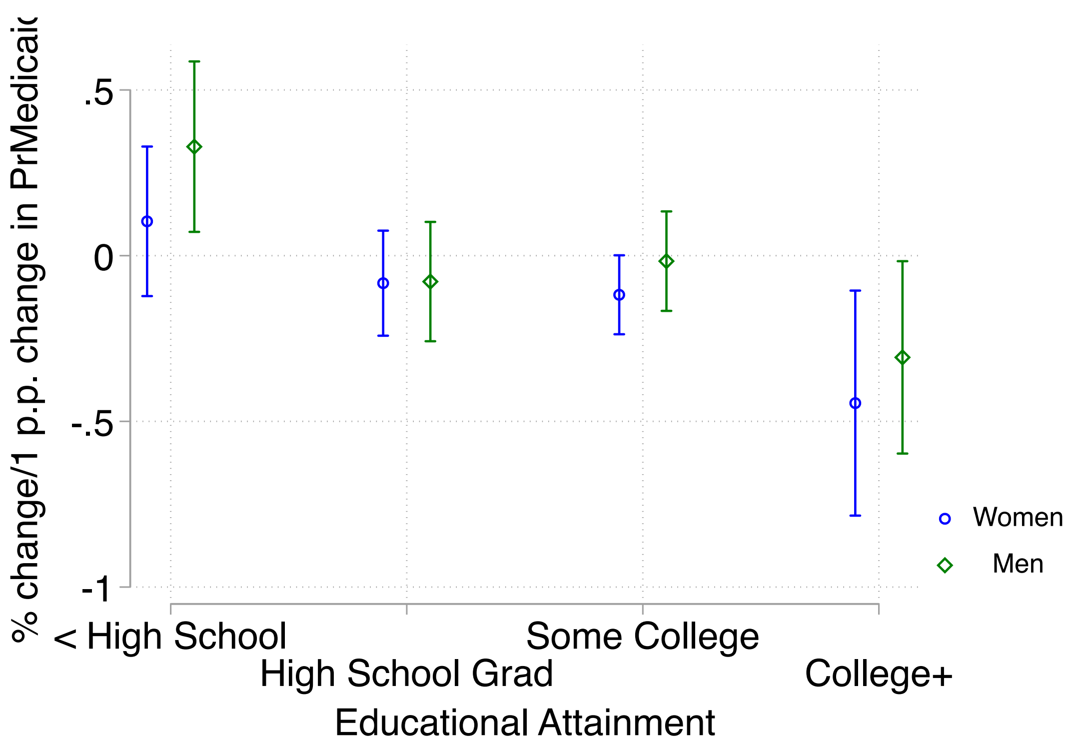


Figure C.4: Effects of Eligibility on Contemporaneous Divorce by Education

(a) Controlling for the fraction of movers



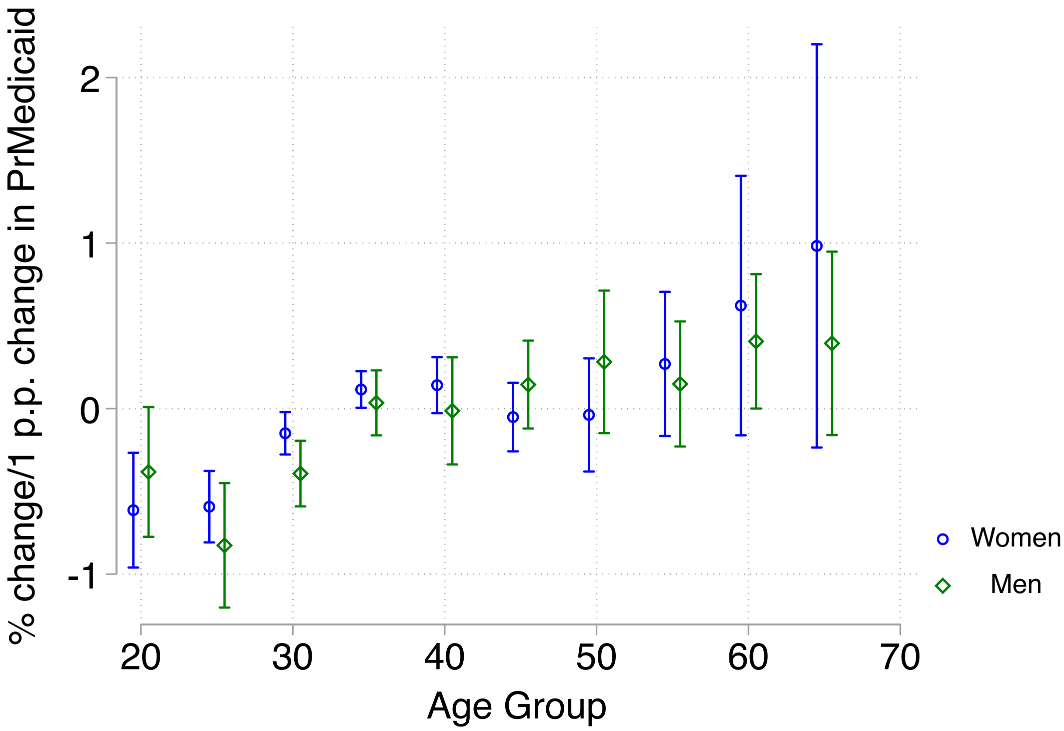
(b) Excluding recent movers



D Addressing potential policy endogeneity

Figure D.5: Effects of Eligibility on Marital Outcomes by Age

(a) Controlling for the fraction of movers



(b) Excluding recent movers

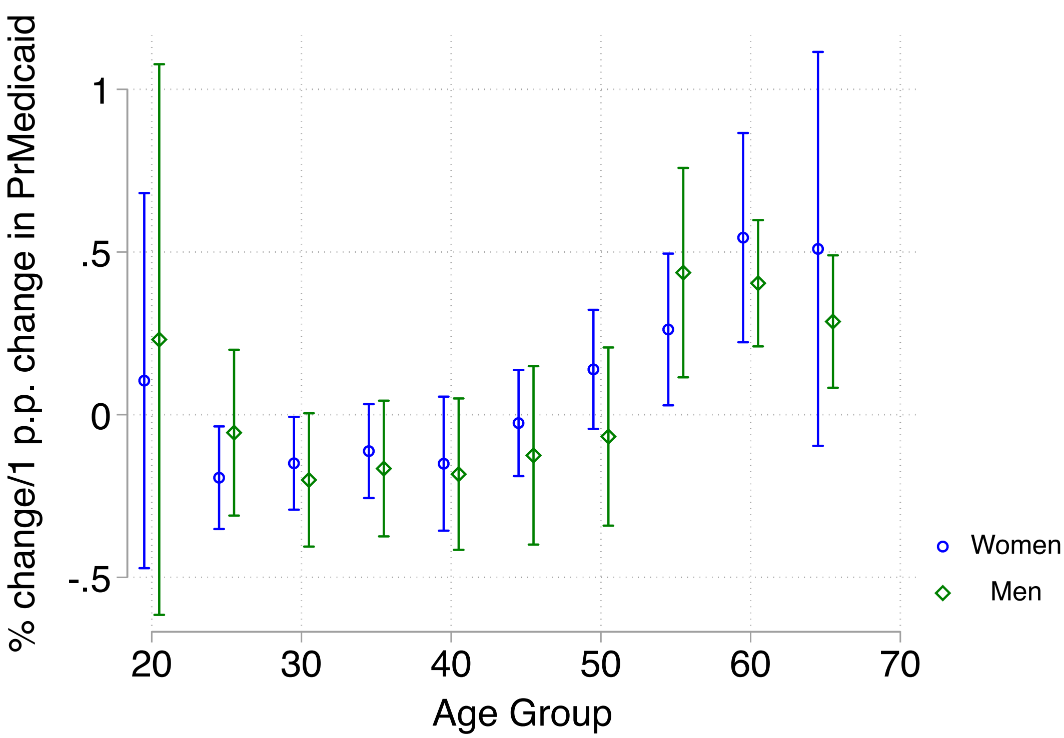
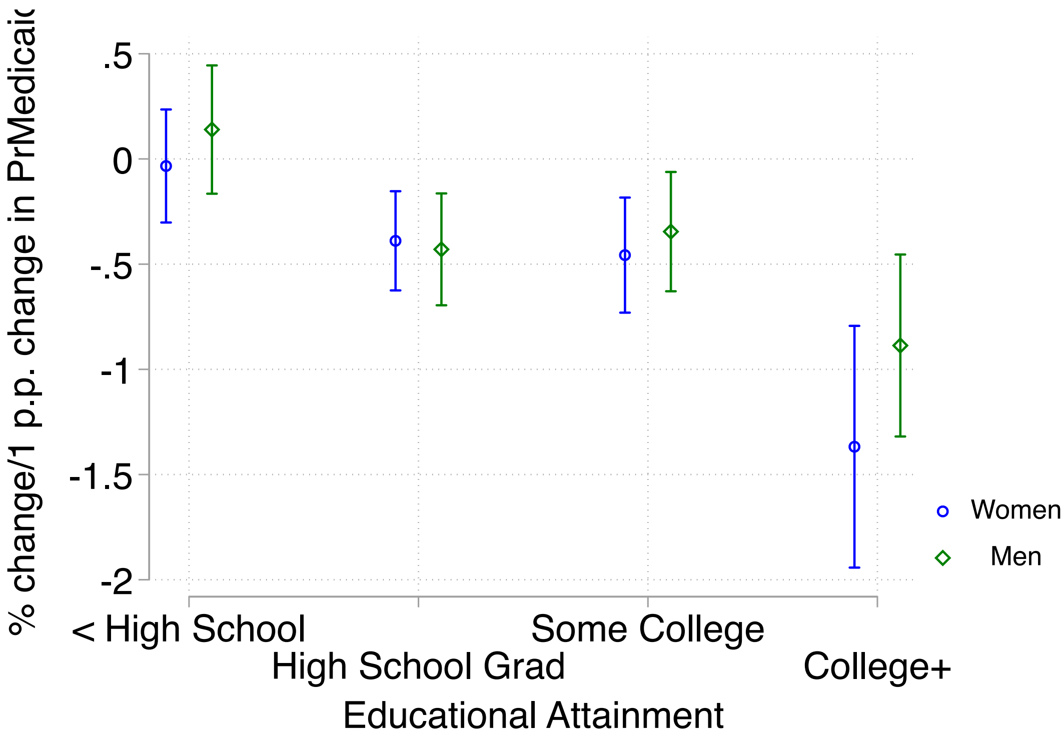


Figure D.6: Effects of Eligibility on Marital Outcomes by Education

(a) Contemporaneous marriage



(b) Contemporaneous divorce

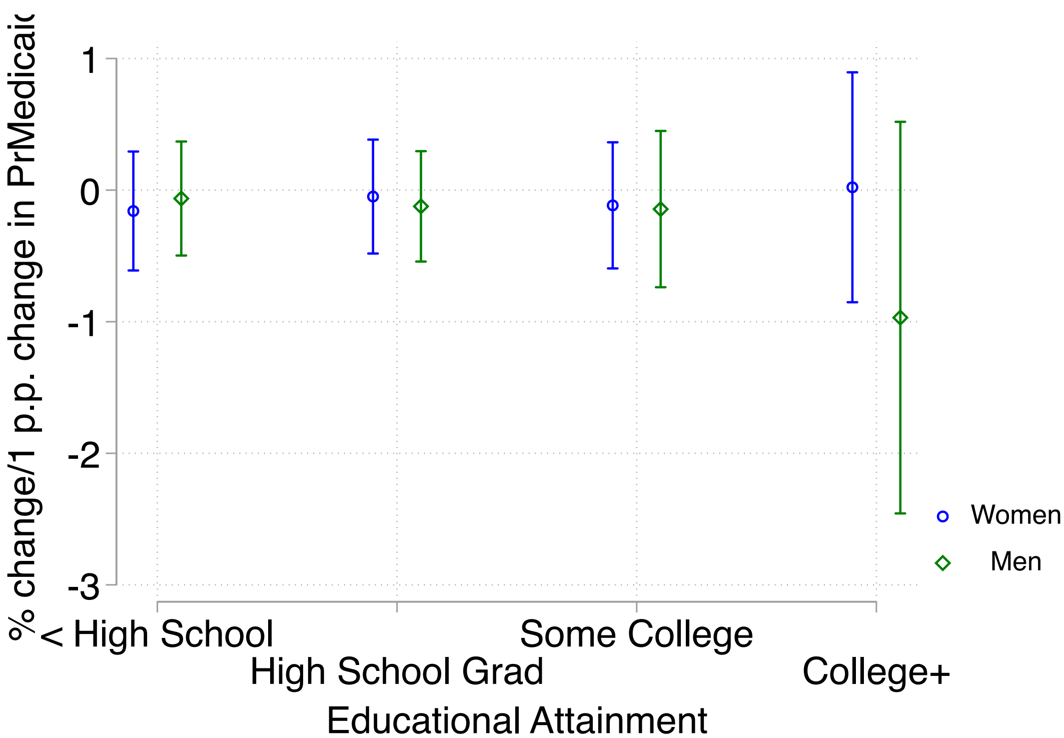


Figure D.7: Effects of Eligibility on Contemporaneous Divorce by Age