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

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Abstract

In addition to love, marriage provides insurance. Public insurance programs for individuals decrease the relative insurance value of marriage. We explore how this tradeoff affects marriage in the context of the Affordable Care Act's Medicaid expansion in the United States. We show theoretically and empirically that while Medicaid expansion does decrease the marriage rate, it also reduces the divorce rate of new marriages, consistent with an increase in match quality. Moreover, it reduces the divorce rate of already-married couples, which we reconcile via the general equilibrium effects on intra-household allocations. Finally, we show that these effects are observed broadly across the population, not only concentrated among individuals with low earnings potential. Taken together, our findings illustrate that even when public insurance reduces the monetary benefits of marriage, the effects on marriage overall may be positive and widespread.

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

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1 Introduction

In the modern romantic view, marriage is about love. Economists, perhaps among the less romantic members of modern society, have mostly ignored this "love" factor and focused on analyzing the considerable economic benefits of marriage. One of the most important economic motives to enter marriage is that it can function as a source of insurance against all sorts of adverse shocks facing individuals. Yet most developed economies also have sizeable public safety net programs, which form an alternative source of insurance for individuals.

In this paper, we explore how the expansion of public insurance for individuals impacts the non-economic benefits of marriage (what non-economists might call 'love') and marital outcomes by reducing the value of intra-household insurance from marriage. To do so, 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').

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

Fig. 1 shows that the gaps in health insurance coverage between married and non-married adults shrink markedly around the year 2014, when many key ACA reforms including Medicaid expansion were largely implemented. Analysis of the ACA's effect on 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.

Perhaps unique among high-income countries, health insurance in the United States is important not only for health but for 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 put households into debt and bankruptcy.[1] Our interest the Medicaid expansion may affect marital outcomes is primarily via this financial channel.

Previous work studying the effects of Medicaid, even in the context of its impact on financial outcomes, have focused mostly on its direct impacts via increased coverage and eligibility. However, in relation to the marriage market, we posit that there may also be substantial indirect effects of Medicaid expansion. In making marital decisions, married individuals may take into account the availability of public insurance programs for non-married individuals; singles may consider the likelihood to become eligible in the event of

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

a negative income shock.

To study the direct and indirect channels, we formulate a stylized search-based equilibrium model of the marriage market, which is based on the model in Anderberg (2007). While simple, the model is informative about the mechanisms through which public insurance might affect the marriage market.

Individuals in the model face unexpected income shocks. They can 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 is match quality, 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: it takes time. 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] With public insurance, the critical match quality level 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 newly formed matches will be higher. If we 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.

In the second half of the paper, we bring these theoretical predictions to data. Our aim is to assess the causal effect on the marriage market of an increase in the probability of Medicaid coverage when not married.

To address the identification problem, we will adopt a 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 given demographic characteristics would be eligible for Medicaid in a certain state and year, as a direct proxy for the unobserved probability to be eligible for Medicaid when non-married.

We use the American Community Survey over the years 2010-2017 to conduct the analysis. We show that an increase in $PrMedicaid_t$ decreases the marriage rates for men and women, and an increase in the probability of eligibility at the time of marriage decreases subsequent divorce rates of newly married couples. While we cannot directly measure match quality, these patterns are consistent with changes we would expect to observe in the marriage market if match quality increased.

We also find evidence that the divorce rate of already-married couples decreases in response to increases in the probability of eligibility for Medicaid outside of marriage, and

[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.

that married couples exhibit shifts in labor supply, both of which are consistent with a shift in intra-household allocations as a result of the general equilibrium effects of public insurance on the marriage market.

Finally, we show that though Medicaid is a means-tested program, we observe responses across a broad range of the population: even highly educated individuals have a lower likelihood of marriage and divorce in response to an increase in the probability of Medicaid coverage when not married.

Our paper makes several contributions. First, our study of the effects of the ACA-era Medicaid expansion provide implications for the effects of public insurance more generally 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.

Second, we use insights from a general equilibrium model of the marriage market to to demonstrate that Medicaid (and public insurance) can affect not only marriage and divorce probabilities, but also match quality and intra-household allocations, which the previous literature on welfare and marriage has not explored. These insights build upon the theoretical model of Anderberg (2007) analyzing the relationship between intra-household risk-sharing and public insurance on marital outcomes. This contribution is similar to work showing that public policy changes that affect the value of marriage relative to singlehood can affect the marriage market equilibrium and intra-household allocations (Chiappori and Oreffice, 2008; Low et al., 2018) as well as match quality (Persson, 2020).

Relative to these two bodies of work, a novel contribution of our work is to highlight that even policy targeted towards the low-income can have effects on marriage quality and intra-household allocations broadly across the population. Persson (2020) and Chiappori and Oreffice (2008) show similar effects of *universal* public policies (respectively, widow's pensions and contraceptive pills), while Low et al. (2018), like much the earlier research on welfare and marriage, focus on the effects of U.S. welfare reform on only women with low educational attainment. Our findings suggests that public safety net programs like Medicaid may have impacts on behavior of individuals even who have a low propensity to currently or ever become eligible for coverage, implying broader impacts than previously recognized.

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

[3]For example, two 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.

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 moniker 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 FPL, while 18 states had eligibility levels over 100% of the federal poverty line (Musumeci, 2012).

Given the patchwork of options for health insurance coverage, 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 unsurprisingly 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 Key Features of the ACA

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]

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. Iin 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. 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

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

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 Theoretical Framework

3.1 The environment

In this section, we present a stylised model that highlights how Medicaid expansion might affect the marriage market. 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 are focused on Medicaid as the specific empirical setting in this paper, the theoretical results are more broadly generalizable to public insurance programs for individuals against 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. To be more precise, we assume that the utility flow for each individual depends on (i) marital status and (ii) the realization of the shock. Hence, individual i receives utility $\psi^i(\omega)$, in case i is married and where $\omega \in \{-1, 1\}$, furthermore we assume $\psi^i(-1) < \psi^i(1)$.[5] Similarly, for a single individual i the utility equals $\psi_0^i(\omega)$ and $\psi_0^i(-1) < \psi_0^i(1)$. The shock $\omega = -1$ indicates an exogenous loss in (indirect) utility. Given the stylized nature of our model, this shock can have multiple sources/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.[6]. The situation where $\omega = 1$ denotes the baseline case. The

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

[6] Given that the focus of our paper is on the effects of Medicaid expansion as a public insurance scheme,

associated probability of receiving a shock $\omega = -1$ is denoted by π.

Match quality and intra-household transfers There are two main reasons why individuals would like to marry in our model: the possibility to obtain private insurance against the contingency of an adverse shock and some non-material benefit, match quality, which can be interpreted as 'love'.

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. [7]

Utilities 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[8], in particular, in 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:

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

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.[9] To be more precise, each single can meet another

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.

[7]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.

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

[9]Given the strong symmetry present in our model, e.g. equal gender ratio, 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.

single of the opposite sex with a probability[10] $\phi(S)$, where $\phi(0) = 0$ and $\phi'(S) > 0$ for all $S \in (0,1)$.

3.2 Equilibrium

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

$$r\mathcal{V}_0^i = \tilde{\psi}_0^i + \phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{V}^i\left(\tilde{\theta}\right) - \mathcal{V}_0^i, 0\right\}\right]. \tag{1}$$

The first term in (1) 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:

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

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 a new match quality $\tilde{\theta}$ with the value outside 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:

$$\begin{aligned} r\mathcal{V}^f(\theta) &= \tilde{\psi}^f + \tilde{t} + \tfrac{1}{2}\theta \\ &+\lambda\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{V}^f\left(\tilde{\theta}\right) - \mathcal{V}^f(\theta), \mathcal{V}_0^f - \mathcal{V}^f(\theta)\right\}\right]. \end{aligned} \tag{3}$$

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, $\mathcal{S}(\theta)$, in particular:

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

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

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

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

This implies the following:

$$\mathcal{V}^m(\theta) = \mathcal{V}_0^m + (1-\gamma)\,\mathcal{S}(\theta), \tag{4}$$

for married men and

$$\mathcal{V}^f(\theta) = \mathcal{V}_0^f + \gamma\mathcal{S}(\theta), \tag{5}$$

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 (1), (2)-(3) and (4)-(5), we can write down the continuation value for the marriage surplus (for a match with match quality level θ):

$$\begin{aligned}(r+\lambda)\,\mathcal{S}(\theta) &= \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f + \theta \\ &+\lambda\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right] \\ &-(1-\gamma)\,\phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right] \\ &-\gamma\phi(S)\,\mathbb{E}_{\pi,G}\left[\max\left\{\mathcal{S}\left(\tilde{\theta}\right), 0\right\}\right].\end{aligned} \tag{6}$$

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[11]:

$$\underline{\theta} + \frac{\lambda}{r+\lambda}\varphi(\underline{\theta}) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f = \frac{\phi}{r+\lambda}\varphi(\underline{\theta}). \tag{7}$$

The first step to solve for an equilibrium is to find a unique solution $\underline{\theta}$ for (7), conditional on the pool of singles, S. Our first result is that this is always possible[12]:

Proposition 3.1 (cut-off match quality)**.** *Given G, λ, π, r, S and ϕ, there exists a unique $\underline{\theta}$ which satisfies:*

$$\underline{\theta} + \frac{\lambda - \phi(S)}{r+\lambda}\varphi(\underline{\theta}) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f = 0. \tag{8}$$

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

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[11] 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 $\varphi(\underline{\theta}) = \mathbb{E}_{\pi}\int_{\underline{\theta}}^{\infty}\left[1-G\left(\tilde{\theta}\right)\right]d\tilde{\theta}$, and substitute this into the equation $\mathcal{S}(\underline{\theta}) = 0$, we obtain (7).

[12] All the proofs are contained in the Appendix.

Given a solution $\underline{\theta}$ to (8), we can find the corresponding size for the pool of singles. Indeed, we will follow **?**anderberg2007marriage) 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:

$$\phi\left(S\right) S\left[1-G\left(\underline{\theta}\right)\right]=\lambda\left(1-S\right) G\left(\underline{\theta}\right). \tag{9}$$

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 (9) indeed states that in steady state and for a given level $\underline{\theta}$ these flows should be equal. Furthermore, (9) implicitly defines[13] a function $\underline{S}=\underline{S}\left(\underline{\theta}\right)$. 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 3.2 (Marriage market equilibrium)**.** *Given the primitives of the model, $\{\psi^i\left(\omega\right), \psi_0^i\left(\omega\right), \pi, G, \lambda, \varphi, r\}$, 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* (8) *and*
2. *The value of $\underline{\theta}$ produces a solution, $\underline{S}$, for* (9) *that yields a match probability consistent with $\underline{\theta}$.*

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 (8) it can be shown that $\underline{\theta}$ is increasing in ϕ and from (9) 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 3.3 (existence of a marriage market equilibrium)**.** *Given the primitives of the model, $\{\psi^i\left(\omega\right), \psi_0^i\left(\omega\right), \pi, G, \lambda, \varphi, r\}$, a marriage market equilibrium $(\underline{\theta}, \underline{S})$ always exists.*

An important remark here is that, even though Proposition 3.3 shows existence of a marriage market equilibrium, it does not imply uniqueness of an equilibrium, as there generally will be multiple equilibria[14]. In case we further assume that the support of match quality is sufficiently large, then it is natural to assume that there exists $-\infty < \theta^L < \theta^U < +\infty$ 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 $S \in (0, 1)$.

[13]It can be shown that (9) defines a unique solution, $\underline{S}$ for a given $\underline{\theta}$.

[14]This is a general feature for search-based equilibria models of the marriage market and it also applies to the frameworks presented in **?**anderberg2007marriage) and Goussé et al. (2017).

3.3 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[15] to help insure individuals against adverse shocks, i.e. the contingency in which $\omega = -1$. The amount of transfer 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. $\omega^m = \omega^f = -1$.[16] 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:

$$\underline{\theta} + \frac{\lambda - \phi}{r + \lambda} \varphi\left(\underline{\theta}\right) + \tilde{\psi}^m + \tilde{\psi}^f - \tilde{\psi}_0^m - \tilde{\psi}_0^f + \pi\left(\pi - 2\right)\tau = 0. \tag{10}$$

The difference between (8) and (10) is the presence of the term $\pi\left(\pi - 2\right)\tau$, which is negative. The condition in (10) implicitly defines a function $\underline{\theta}\left(\tau\right)$.We can then easily show the following result:

Lemma 3.4 (New cut-offmatch quality). *Assume $\lambda \leq \phi\left(S\right)$, then $\underline{\theta}\left(\tau\right)$ 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 3.4 is very intuitive: the public transfer scheme offers a substitute to one of the main benefits of being in a partnership, that is, insuring each other against shocks. Hence, in order to still enter or stay in marriage, the non-material benefit (i.e. match quality) has to increase.

[15] 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, theyassume Medicaid operates as a consumption floor, i.e. a minimum level of consumption which the government guarantees to all individuals.

[16] 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. hence, eligibility of individuals depends fundamentally on the characteristics of the household overall and not only his/her circumstances. Hence, 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.

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[17],denoted by $\underline{S}^{L}(\tau)$ and $\underline{S}^{U}(\tau)$ respectively. To make progress, we first observe that the result in Lemma 3.4, together with the monotonicity (and continuity) of $\underline{S}(\underline{\theta}(\phi(.)))$, 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 3.5 (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 measure of singles 'upwards'.

3.4 Extensions

Though the framework presented so far is relatively stylised and omitted several aspects which might be of importance, it already provided 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.

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, **?**anderberg2007marriage) 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.[18] 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 $\lambda G(\underline{\theta})$. Hence, 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

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

[18] 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).

in relatively higher match quality matches, which then remain of 'high quality' for some amount of time due to the persistence in match quality.

Intra-household allocations.

The couple was assumed to split the marital surplus through generalised Nash bargaining, as highlighted by equations (4) and (5). 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,

$$\tilde{t} = (1-\gamma)\, r\mathcal{V}_0^f - \gamma r\mathcal{V}_0^m + \gamma\tilde{\psi}^m - (1-\gamma)\,\tilde{\psi}^f + \left(\gamma - \frac{1}{2}\right)\theta. \tag{11}$$

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, hence, 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)).[19] Assuming now $\tau^f > \tau^m$, then the outside value for women increases more than for men. From (11) it is clear this alternative reform increases the transfer from the husband to the wife. In the absence of explicit intra-household decisions (e.g. labor supply, consumption, household production, childcare decisions), this is the only 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: males versus females), the public transfers associated with the insurance will have broader effects in terms of allocations within the household.

Notice that these results are similar to what one would obtain in a limited commitment collective household framework[20] in which it is assumed that spouses choose intra-household allocations (in terms of consumption, time use decisions, and so on) which are ex-post Pareto optimal. Building on the work on recursive contracts (e.g. Marcet and Marimon

[19] 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, such as 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.

[20] For a recent overview on intertemporal collective household models, we refer the interested reader to Chiappori and Mazzocco (2017).

(1992)), one can show that this is equivalent to maximizing a weighted sum of utilities, in which the weights are interpreted as the bargaining power for that particular spouse.[21]

Limited commitment implies that spouses are able to revise their relative bargaining power at any point in time, which will be the case as soon as their participation constraint becomes binding (that is, the value of staying inside the marriage is lower than the outside option of leaving the marriage and searching for a new partner). If such a revision in bargaining power where both spouses are at least as well off inside marriage as outside is no longer possible, the marriage will dissolve. In the present setting, such a limited commitment collective household framework would imply that the public insurance scheme increases the outside value and therefore could make the participation constraint of one spouse binding. The latter spouse would then see an increase in his/her bargaining power, leading to a higher consumption share or higher amount of leisure (hence reduced labor supply).

3.5 Taking stock of the theoretical results

Before finishing the theory section, it is useful to summarize the results we derived from the model, in such a way that we can then bring these theoretical predictions to the data and test whether we can find evidence in favor of the search-based equilibrium model of the marriage market. In particular we derived the following testable 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.

> **Theoretical Prediction 2: change in intra-household allocations.** 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.

4 Empirical Strategy

The goal of the empirical strategy is to assess the predictions of the theoretical framework regarding the effect of public insurance on marital outcomes. We study these predictions in the context of the Medicaid expansions that followed passage of the Patient Protection and Affordable Care Act in the United States.

[21] Many papers have exploited the mathematical structure analyzed by Marcet and Marimon (1992), including Kocherlakota (1996), Ligon (1998), and Ligon et al. (2002) who use this structure to analyze dynamic risk sharing and papers studying intertemporal household allocations, such as Mazzocco et al. (2007), Voena (2015) and Lise and Yamada (2018).

4.1 Identification

We aim to identify the causal effect of an increase in the probability of becoming eligible for Medicaid when non-married on marital outcomes. To do so, we employ a strategy similar in spirit to the simulated instruments approach pioneered by Currie and Gruber (1996a) and Currie and Gruber (1996b). The motivation for its use in our setting is distinct from how it has been employed in previous studies of Medicaid expansion, however: typically the true treatment of interest has been individual eligibility or individual coverage through Medicaid, and the instrument is constructed to account for the endogeneity of individual eligibility or coverage in a two-stage estimation strategy. The instrument captures variation in the likelihood of eligibility that is plausibly due to state-level policy changes and uncorrelated with individual characteristics.

In our setting, we do not observe Medicaid eligibility when single for currently-married individuals. Thus, we use a simulated instrument to proxy for the probability of being eligible when single.

There are two components to this measure. First are the eligibility rules for Medicaid, which indicates a threshold $\bar{y}$ for which an individual of type j would be eligible for Medicaid in a given state and year. The second component is the probability that an unmarried individual of type j would receive an income of $\bar{y} < y$. We capture this probability by calculating the fraction of individuals of type j that earn less than this amount in a national fixed sample, so the underlying earnings distribution from which these fractions are calculated is common to type j across states and years. These earnings distributions allow us to characterize the likelihood that an individual's income would fall below any particular level (analogous to the notion of π in the theoretical framework). By combining these two components, we construct a measure of the probability that individuals of type j would be eligible for Medicaid in a given state and year.

The key identifying assumption is that this measure of eligibility for Medicaid is unconfounded by other factors that may affect marital outcomes. Before discussing potential threats to identification, we will describe the data we use in the analysis.

4.1.1 Data and variable construction

We use data from the American Community Survey between 2010 and 2017, 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.[22] 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,

[22]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

family size, and the state in which they reside.[23]

To determine eligibility rules by state and year, we use information gathered from yearly reports between 2010 and 2017 by the Kaiser Commission on Medicaid and the Uninsured (see Appendix A 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.[24]

To calculate the key variable of interest, the probability that an individual would become eligible for Medicaid when 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, race (white, African-American, and other races, are grouped due to their small sample size), 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.

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. In addition to basic demographic characteristics in Panel (a), Panel (b) reports information on health insurance status and marital transitions. The vast majority of individuals over the whole sample have some health insurance coverage, though fewer men than women (a coverage rate of 83% for men and 87% for women), keeping in mind that this statistic includes both pre- and post-Obamacare implementation. Of these, 11% of men and 14% of women report coverage by Medicaid. The estimated probability of eligibility for Medicaid when not married is slightly lower than the observed coverage of Medicaid for men at 10%, and higher for women at 16%, which likely owes to that women are more likely to become eligible when not married (since any children usually are counted in their household) while men are slightly less likely to become eligible when not married.

[23]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.

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

Panel (c) provides basic statistics on marital status and transitions. About half of the sample of adults are married across all years, and about 10% are divorced at any given time (this does not define "ever divorced", simply those who have not remarried). The average annual rate of marriage is approximately 3% and the rate of divorce is about 1.5%. The flow into marriage is calculated by taking the fraction of individuals who report being married in the past year relative to all individuals who were not married one year ago (which is the sum of those who report being married in the last year and those who report being currently non-married with no change in marital status in the last year).

In one part of the analysis, we use only individuals who were married between 2010 and 2017, and Table 2 provides summary statistics on this group in particular. Unsurprisingly, they are on average younger than the full sample, more highly educated, with a larger education differential by gender, and have fewer children. Likely because they are more highly educated than the full sample, they also have lower probabilities of becoming eligible for Medicaid when not married.

To give some intuition for the identifying variation exploited, Fig. 2 graphs $PrMedicaid_{d,s,t}$ for two demographic groups, showing the estimated eligibility for six states for both parents and childless individuals who are white, 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. Finally, Wisconsin is a state that has ostensibly never participated in Medicaid expansion, yet in 2014 both inreased 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.

4.1.2 Threats to identification

There are two primary possible issues with the identifying assumption of exogeneity: state economic or political conditions and potentially endogenous migration.

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. During the period that we are studying, in practice most of the variation in state eligibility is in response to the federally legislated rules of the ACA.

However, similar to non-ACA related changes in eligibility rules, selection into Medicaid expansion itself could potentially be correlated with local political or economic conditions, since expansion was left up to the states. To address this concern, in robustness checks we add controls for whether the state had a Democratic, Republican, or split government[25] as well as controls for local economic conditions, to account for the possibility that states with higher need among the population may have had greater incentives to accept the subsidized federal expansion.

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.

4.2 Estimation

In this section, we describe how we aim to test empirical analogues to the three theoretical predictions given in Sec. 3 regarding marital match quality, intra-household bargaining in the new marriage market equilibrium, and behavioral responses to the indirect benefits of insurance.

4.2.1 Empirically testing for the effect on marital match quality

Marital match quality does not have a straightforward empirical analogue. While we can characterize it abstractly as the "non-pecuniary" benefits of marriage, or more specifically but no more concretely as "love", it is not obvious whether or how one might measure match quality quantitatively. Rather than attempting to measure it directly, then, we test two empirical predictions that we would expect to observe in relation to marital transitions if the quality of marital matches in the new equilibrium had improved as a result of expanded Medicaid eligibility for non-married individuals.

Empirical Prediction 1: Decreased marriage rate. As non-married individuals' outside options improve, they become pickier about spouses, and the minimum quality of

[25] In other words, whether the governor and houses of the state legislature were all of one party or whether they were split.

a match for which they will be willing to marry rises. As a result, we expect to observe relatively fewer transitions from being single to married where Medicaid eligibility is higher.

The equation that we will estimate associated with this prediction is:

$$\text{MarrLastYear}_{d,s,t} = \eta \text{PrMedicaid}_{d,s,t} + \mathbf{X}_{d,s,t}\beta + \varepsilon_{d,s,t} \tag{12}$$

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 $PrMedicaid_{d,s,t}$, which characterizes the probability that an individual would be eligible for Medicaid outside of marriage, and $X_{i,s,t}$ represents a vector of control variables. The vector of control variables includes full sets of dummies for year and state and each of the demographic variables (age group, race, educational attainment). Additional specifications interact the demographic variables with the full set of year dummies to allow for flexible time trends, as well as including state-specific control variables for local economic conditions and measures of political control. Standard errors are clustered at the level at which eligibility legally varies, which by state, year, and parental status.

Empirical Prediction 2: Decreased incidence of divorce for higher eligibility at time of marriage. The theoretical model predicts that as long as the distribution of shocks to match quality does not change, the increase in minimum match quality in response to a higher level of public insurance eligibility at the time of marriage will result in a lower propensity to divorce in any subsequent time period. Empirically, this translates to a prediction that couples married in states and years with higher eligibility should have lower incidence of divorce in a given year of marriage than similar couples married under lower eligibility likelihoods.

To test this prediction, we instead assign $PrMedicaid_{d,s,t^*}$ 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 eligiblity 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. **finish

We associate $PrMedicaid_{d,s,t^*}$ given these inferred characteristics at the time of marriage. The eligiblity levels are otherwise calculated exactly as before.

We estimate the following equation in relation to this prediction:

$$\text{DivLastYear}_{d,s,t} = \eta \text{PrMedicaid}_{d,s,t^*} + \mathbf{X}_{d,s,t^*}\beta + \gamma_{t-t^*}\varepsilon_{d,s,t} \tag{13}$$

The structure of the equation is similar to Eq. 12 although the right-hand side variables are with respect to characteristics at the time of marriage, as well as dummy variables for the years elapsed since marriage to account for common trends in the likelihood of divorce over the course of a marriage.[26]

4.2.2 Empirically testing for the effect on intra-household bargaining power

We also predict that the expansion of a public insurance program like Medicaid will have impacts on intra-household bargaining power of already-married couples, when taking into account that the value of singlehood is likely to increase more for women than for men as a result of the expansion. We consider that a change in intra-household bargaining may affect both the stability of existing matches as well as decisions made within the household.

Empirical Prediction 3: Divorce rates for already-married couples. We predict that Medicaid expansion will affect the divorce probability of already-married couples.

One mechanism by which it might have an impact is quite intuitive: if the value of singlehood increases, married individuals have an increased incentive to become divorced. This aspect would tend to increase the divorce rate of already married couples. Indeed, if no renegotiation is possible in within the household (such as if couples' behavior was represented by a unitary model of the household), we would expect a strictly non-negative impact of Medicaid expansion on divorce.

[26]After controlling for dummy variables 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.

A substantial literature has attested to the fact that couples do not in fact behave as in a unitary model (see Chiappori and Mazzocco (2017) for a review of this work). Couples can renegotiate intra-household allocations (captured in the model by the transfers made within the household) when the participation constraint of one member of the couple binds. This feature usefully captures the intuition that couples are likely not in a constant state of negotiation in response to every shock in the marriage market environment, but that renegotiation likely does occur from time to time. If during such a period of renegotiation, the couple cannot achieve a new allocation such that both members prefer to stay in marriage, the couple divorces.

In our model, some fraction of couples in each time period are experiencing shocks to the quality of their relationship that may induce this type of renegotiation. A simultaneous change in the outside option such as by the expansion of Medicaid are predicted to affect the outcomes of such renegotiations. However, even if we assume that the expansion of Medicaid would disproportionately benefit women, it's not possible to predict whether the impact via its effect on renegotiation would be to increase or decrease divorce rates.

Although the expansion of Medicaid is expected to decrease the surplus of marriage (since the outside options for both men and women improve), the impact of renegotiation may be to make a less-satisfied partner better off without necessarily making the more-satisfied partner substantially worse off, in which case their marriage may become more robust to negative shocks to the quality of their relationship. This is the subtle way in which divorce rates might actually decrease for already-married couples.

To assess this prediction, we estimate the following equation:

$$\text{DivLastYear}_{d,s,t} = \eta \text{PrMedicaid}_{d,s,t} + \mathbf{X}_{d,s,t}\beta + \varepsilon_{d,s,t} \tag{14}$$

This equation bears similarities to both Eq. 12 and Eq. 13. The right-hand side is identical to Eq. 12, 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. 13 is that we estimate the contemporaneous effect of contemporaneous effect of a change in eligibility on couples that were already married when a change went into effect.

While we use $\text{PrMedicaid}_{d,s,t}$ throughout as a proxy for the probability of eligiblity as a non-married individual, there is a potential problem of interpretation in this setting. While married couples may be affected by changes in eligibility for non-married individuals, as we have posited above, they may also be "directly" affected by increases in their eligibility for Medicaid as a married couple. For example, if a married couple actually gains coverage via Medicaid, this may reduce their vulnerability to financial shocks (as described in Sec. 2.3) and hence marital conflict. Even if they do not directly gain coverage, the knowledge that such a safety net exists may improve peace of mind and reduce marital strife.

In this setting, it's not straightforward to distinguish between these two effects. In principle, one could construct a similar measure for the probability that an individual would become eligible for Medicaid while married, however the challenge is that for non-married individuals in this sample (namely, the divorced) we do not observe the characteristics of their (former) spouse. In practice, then, we would have to infer implicitly that divorced individuals had spouses similar to the still-married people with similar characteristics, which may be problematic given our predicted simultaneous changes in the marriage market itself.

However, we can also gain insight on whether the renegotiation channel or the "direct" channel are relevant by studying the behavior of married individuals only, in which case we observe the characteristics of spouses. We describe such an approach next.

Empirical Prediction 4: Labor supply of married individuals. If Medicaid expansion does shift intra-household bargaining power, we are likely to observe changes in the behavior of married couples. One way of testing this hypothesis is to look at the impact of Medicaid expansion on labor supply, treating it as the complement of leisure. [Will add analysis to look at total work as well, including household production] If leisure is a normal good, then an increase in bargaining power within the household is likely to have the impact of an income effect, increasing leisure or alternately decreasing labor supply.

In this case, we can also control for whether the couple is likely to gain Medicaid coverage, in addition to individually accounting for each of the spouse's likelihood of becoming eligible when not married. If the main effect of Medicaid expansion goes via the "direct effect"–in other words, if the measure of the probability of becoming eligible when non-married is primarily capturing the increased likelihood of becoming eligible *when married*, then we may expect that the signs of the variables associated to each spouse should be the same in relation to the impact on the labor supply of wives or husbands respectively. In contrast, if they capture the effect of the changes on the spouse's outside options, these variables may have opposite signs in relation to labor supply.

For this prediction, we estimate the following equation:

$$\begin{aligned}\text{AnnHours}^j_{d,s,t} = {} & \eta_w \text{PrMedicaid}^w_{d_w,s,t} + \eta_h \text{PrMedicaid}^h_{d_h,s,t} + \\ & \eta_{Marr} \text{PrMedicaid}^M arr_{d_w,d_h,s,t} + \mathbf{X}_{d_w,s,t}\beta_w + \\ & \mathbf{X}_{d_h,s,t}\beta_h + \varepsilon_{d_w,d_m,s,t}\end{aligned} \tag{15}$$

The subscripts w and h stand for wives and husbands respectively, and the superscript j indicates whether the outcome variable is the average hours for wives or husbands. Thus, this equation separately estimates the effect of wives' and husbands' potential eligibility for Medicaid when not married based on each of their demographic characteristics. Additionally, it includes a control for the married couple's probability of becoming eligible for Medicaid, which is calculated analogously to that for non-married individuals, except it uses the set of married individuals from the fixed national sample and is based upon the characyeristics of both the husband and wife.

5 Results

5.1 Marital Outcomes

Table 3 presents results from estimating Eq. 12 for men and women. Cols. 1 and 3 include the basic controls described in the estimation, and Cols. 2 and 4 also include interactions of each of the demographic variables with a full set of year dummies to allow for flexible time trends by group. For both men and women, the estimated effect of an increase in $PrMedicaid$ on the marriage rate is negative and highly significant. Considering the specification with flexible time trends, the results imply a 10 p.p. increase in the probability of eligibility for Medicaid leads to a 0.3 p.p. decrease in the marriage rate for men and a 0.65 p.p. decrease in the marriage rate for women. In terms of effect size, these are quite large changes: compared to the sample annual average marriage rate of approximately 3%, these represent roughly a 10% decrease for men and a 20% decrease for women.

Table 4 shows results from estimating Eq. 13, examining whether higher eligibility levels at the time of marriage affect subsequent divorce rates. Though it has a similar structure to Table 3, it's important to keep in mind that the right-hand variables are defined (or inferred) with respect to the time of marriage. Across specifications for both men and women, the point estimates are quite similar, implying that a 10 p.p. increase in eligibility at the time of marriage decreases divorce rates in the first years of marriage by between 0.06 p.p. and 0.08 p.p. While these numbers may seem small in absolute value, given an approximate annual average divorce rate of 1.5%, such a change represents nearly a 5% reduction in the divorce rate. Moreover, many demographic groups (particularly childless individuals) were subject to substantially larger changes in the probability of eligibility.

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. 3 displays the coefficient on $PrMedicaid$ when estimating Eq. 13 separately by years since marriage for men and women. While the estimates are less precise than when pooled as in Table 4 (several of the estimated coefficients are not statistically distinguishable from zero at the 5% level), the patterns for both men and women are suggestive that the point estimates are becoming more negative with years since marriage

5.2 Intra-household Effects

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 5presents results pertaining to this hypothesis, again displaying specifications including the baseline controls (Cols. 1 and 3) as well as with flexible time trends (Cols. 2 and 4). Across specifications and for both men and women, the effect on divorce rates is again negative and significant, although the magnitude of the effect appears to be larger for men than for women. For men, a 10 p.p. increase in the probability of Medicaid eligibility when not married is associated with a decrease of 0.16

p.p.while for women it is only 0.08 p.p.

As discussed in Sec. 4.2, although we find a significant and negative relationship here it is possible that this negative relationship is essentially a function of omitted variable bias, and the true effect comes from the "direct" impact of Medicaid eligibility or coverage *for married couples* rather than changes to their married individuals' potential outside options as singles. However, this is difficult to test with respect to divorce rates without making additional strong assumptions about the characteristics of former partners of the divorced individuals.

We turn instead to an analysis of labor supply where we focus solely on a sample of married individuals, so that we observe the characteristics of each individual's spouse. Table 6 presents results from estimating the reduced-form relationship given by Eq. 15 and variations. Cols. 1 and 4 include only $PrMedicaid$ as defined in all previous regressions, the probability that men (women) of a given demographic group would become eligible for Medicaid when not married. Cols. 2 and 5 subsquently add the equivalent variable for the spouse of a given individual. Finally, Cols. 3 and 6 add a similarly constructed variable for the probability that the married couple themselves would become eligible.

For men, neither the own or spousal probabilities of Medicaid coverage as singles have a significant effect on annual hours. Interestingly, however, the probability that the couple would become eligible does have a significant negative coefficient of 59.8, consistent with the standard hypothesis in the study of Medicaid on labor supply that actual coverage or eligibility may decrease labor supply.

For women, Col. 3 gives a significant negative effect of the probability of own Medicaid coverage on annual hours with a coefficient of -114.1. In contrast, when adding husbands' probability of eligibility, the associated coefficient is actually positive and nearly as large in magnitude, at 78.7, while the coefficient on wives' eligibility instead becomes more negative at -157.9. Adding the control for the couple's probability of eligibility together hardly affects the point estimates. While the implied effects on labor supply are not particularly economically significant, the fact that the nevertheless have significant estimated effects on women's hours of opposite sign suggest that they are not merely capturing the effect of the probability that the couple themselves would be covered, since both spouse's variables are positively correlated with the couple's probability of being eligible.

5.2.1 Effects across Subgroups

In our baseline analysis, we find effects of the probability of Medicaid eligibility on marriage and divorce for the population as a whole. However, Medicaid is targeted towards individuals with low income. In the subgroup analysis, we are interested in exploring whether the average effects for the whole population derive from the lower-income group primarily targeted or whether individuals from the broader population also respond to the policy change.

To do so, we disaggregate the effects by education, interacting the probability of Medicald eligibility with the four levels of education (less than high school; high school graduate; some college; college graduate). Although in a given state, all individual are subject to the same Medicaid rules, those with higher levels of education (and hence expected income) have a lower absolute probability of becoming eligible by Medicaid's income thresholds. As a result, for the same increase in eligibility levels, the higher-educated experience a smaller absolute increase in the probability of becoming eligible for Medicaid coverage than lower-educated individuals.

Fig. 4 shows both the marginal effects of the probability of Medicaid coverage on marriage rates by education, and these effects when scaled by the mean increase in the probability of becoming eligible for states that expanded in 2014 for different educational groups. When expressed in terms of the marginal effect of an increase in the probability of Medicaid coverage as in Panel (a) (where one unit is interpreted as a 100 percentage point increase), the magnitude of the marginal effects increase monotonically with education, and appear to be an order of magnitude larger for college-educated as compared to those with less than a high school degree. However, the same increase in a Medicaid threshold level does not translate to the same increase in probability of coverage across educational levels: since the higher-educated have a higher expected income, the same increase in the threshold yields a smaller increase in the probability of Medicaid eligibility than for the lower educated.

For example, when considering only states that participated in the Obamacare expansion in 2014, the average increase in the probability of Medicaid eligibility for women with less than a high school education was 32 percentage points compared to 8.2 percentage points for college-educated women.

When adjusting for the fact that the absolute change in the likelihood of eligibility for a given reform is smaller for the higher-educated, we see that the differences across educational level are much smaller than those implied by the marginal effects. Panel (b) reflects this fact by scaling the marginal effects by the mean increase in the probability of Medicaid eligibility for a given sex and educational group for states that implemented the Obamacare expansion in 2014.

Both of these depictions of the effects by educational status illustrate that there are substantial behavioral responses among the higher-educated, even though they still have low absolute probabilities of becoming eligible for Medicaid (about 14% of college-educated non-married individuals are eligible under the 138% threshold, in comparison to roughly half of those who have not graduated high school).

Fig. 5 presents a similar pattern. While the effects are not monotonic with respect to education (for women, the effect for high school graduates and those with some college is close to zero and statistically insignificant), for both men and women the effect for college-educated individuals given in Panel (a) appears substantially larger in magnitude. In contrast, the scaled effects in Panel (b) illustrate that for an increase akin to the average increase of 2014 expansion states, the marginal effect for lower and higher-educated was actually similar in magnitude.

We consider a similar breakdown by education for eligibility at the time of marriage, though the results do not lend themselves as easily to interpretation. For women, the marginal effects by education are not statistically significant; perhaps like for the disaggregation by years since marriage there are too few observations per group in this sample. For men, there are noticeable differences by education, although in this case the effect on divorce for the lower-educated is negative while for college-graduate men it is actually positive. Without speculating too much about the reason for the difference in sign, this result still highlights that the highly educated individuals' marital outcomes do appear to be affected by increases in the probability of Medicaid eligibility, whereas it would not have been unreasonable to guess that they should be unaffected or at least in smaller measure than those with lower-expected earnings as proxied by education.

These results highlight that Medicaid appears to affect marital decisions not only the lower-educated, who have a higher likelihood of being low-income, but also those who have a low absolute likelihood of becoming eligible for Medicaid.

5.3 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.

5.3.1 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 is such individuals happen to have lower marriage and divorce rates, as 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.

The ACS data for the time period we study includes information on whether individuals have moved from another state in the past year. Using this, we construct a variable for the fraction of the demographic group that has moved into the current state in the past year. We can use this variable to better understand the potential issue in a few different ways.

First, we consider the fraction of migrants as the dependent variable of a regression equation that is otherwise identical to the right-hand sides in Eq. 12 and Eq. 14. The results for this exercise are presented in Cols. 1 and 2 of Table 7, for men and women respectively, when individuals of all marital statuses are pooled, while Cols. 3 and 4 present results for married men and women (corresponding to the sample used for the labor supply analysis). Contrary to the hypothetical concern, for women eligibility is negatively correlated with the fraction of individuals moving while for men there is no significant relationship.

While it's not obvious why there would be a negative relationship with eligibility, given that there is a significant association for women, it may still be problematic for our identification via another channel that we have not imagined. Given a negative relationship with eligibility for women, it would be particularly concerning if "movers" are also more likely to become married and more likely to become divorced. In the next step, we re-estimate the main results on marriage outcomes while controlling for the fraction of the demographic group that has moved into the state in the past year.[27] These results are presented in Table 8. For the marriage rate, in Cols. 1 and 2, the coefficient on the fraction of movers is significant, large, and positive, but its inclusion has essentially no effect on the coefficient on $PrMedicaid$. For the effect of eligibility at the time of marriage on the subsequent divorce rate, in Cols. 3 and 4, the coefficient on the same variable is not significant and the point estimates are close to zero. For the effect of eligibility on divorce rates of already-married couples, the coefficient is negative and significant for both men and women, but again it does not meaningfully affect the coefficients on $PrMedicaid$.

Though it would appear that migration is related to marriage and divorce rates, these results suggest it does not operate as a confounder for the relationship between the probability of Medicaid eligibility and marital transitions.

5.3.2 Local Economic and Political Conditions

Another potential concern is that 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. The former point we can assess by controlling for additional economic conditions, for which we will include the level of unemployment and the mean log earnings for the demographic group. For the latter point, we test whether the inclusion of controls for partisan control of government affects the results.[28]

Cols. 1-2 and Cols. 5-6 of Tables **??** add the additional controls for men and women respectively the unemployment rate and the mean log earnings within a given demographic group to the baseline specification, while Cols. 3 and 7 add the dummies for Democratic or Republican control of government, and Cols. 4 and 8 include both sets of controls.

The inclusion of these controls do not appear to meaningfully affect the coefficients on $Pr(Medicaid)$, although the coefficients on the unemployment rate and mean log earnings are themselves often significant in relation to marriage and divorce rates. For example, higher unemployment rates are associated with higher divorce rates of already-married couples, and higher average earnings for men but not for women are negatively associated with divorce rates. In contrast, the party control of government appears to have no relation to marriage or divorce outcomes, with estimated coefficients on both Democratic and Republican control generally insignificant and an order of magnitude smaller than the coefficients on $PrMedicaid_t$. These results suggest that even if there may be partisan selection into expansion, it does not appear to be a confounding factor.

[27] When considering the effect of eligibility at the time of marriage, the migration variable corresponds to the year of marriage as well.

[28] The party of the governor in each year and state is based on Ballotpedia's tabulation through 2014. Additional years were manually coded based on election results documented at the U.S. Elections Atlas. Control of the houses of state legislatures comes from the Correlates of State Policy Project, but were only available through 2016.

6 Conclusion

In this paper we shown that public insurance can have major and sometimes counterintuitive impacts on marital outcomes. Within the context of Medicaid expansion, we show that an increase in the probability of Medicaid eligibility leads to a decrease in the marriage rate and also 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 theoretical prediction of an increase in match quality in the marriage market in response to the introduction of expansion of public insurance that benefits singles. We also provide evidence that the increase in probability of Medicaid eligibility leads to a decrease in the divorce rate of married couples and a shift in intra-household labor supply that are both consistent with a shift in intra-household allocations, as would be predicted by general equilibrium effects in the marriage market if public insurance is more valuable for women than men.

Beyond demonstrating the substantial impacts of Medicaid on marital outcomes, an important contribution of our work is to show expansions of public insurance programs targeted at the low-income may still affect the behavior of individuals who may have low absolute likelihoods of ever gaining coverage. This finding suggests that the impact of public safety nets on behavior (and in particular, marital behavior) may be much greater than previously recognized, and could be a useful avenue of further research.

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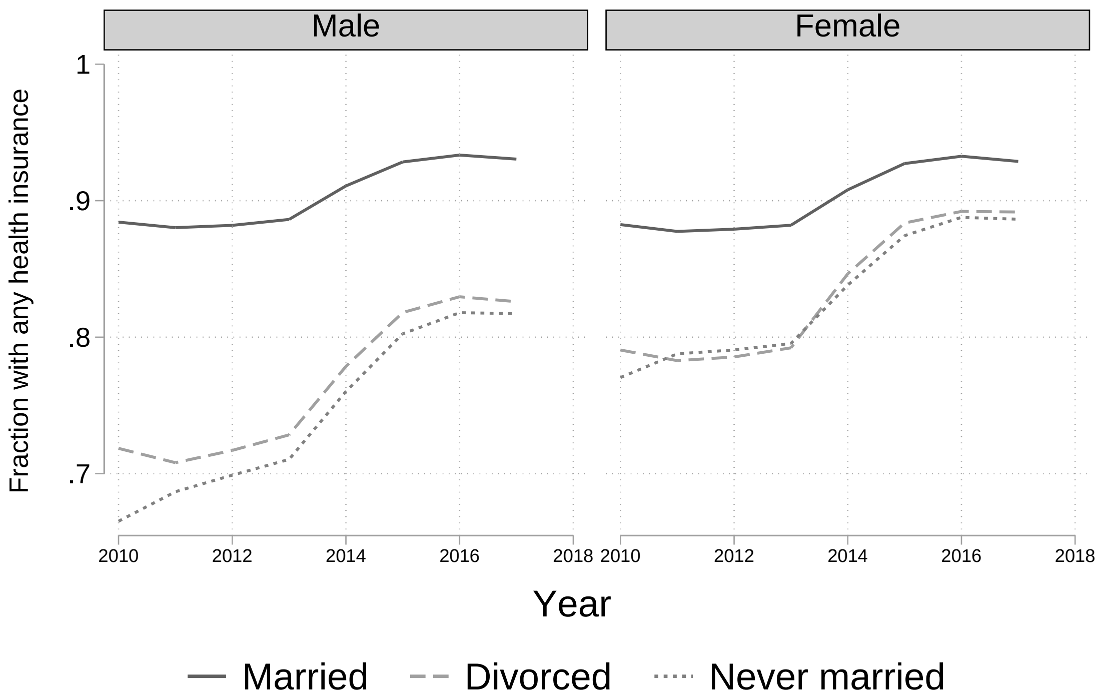
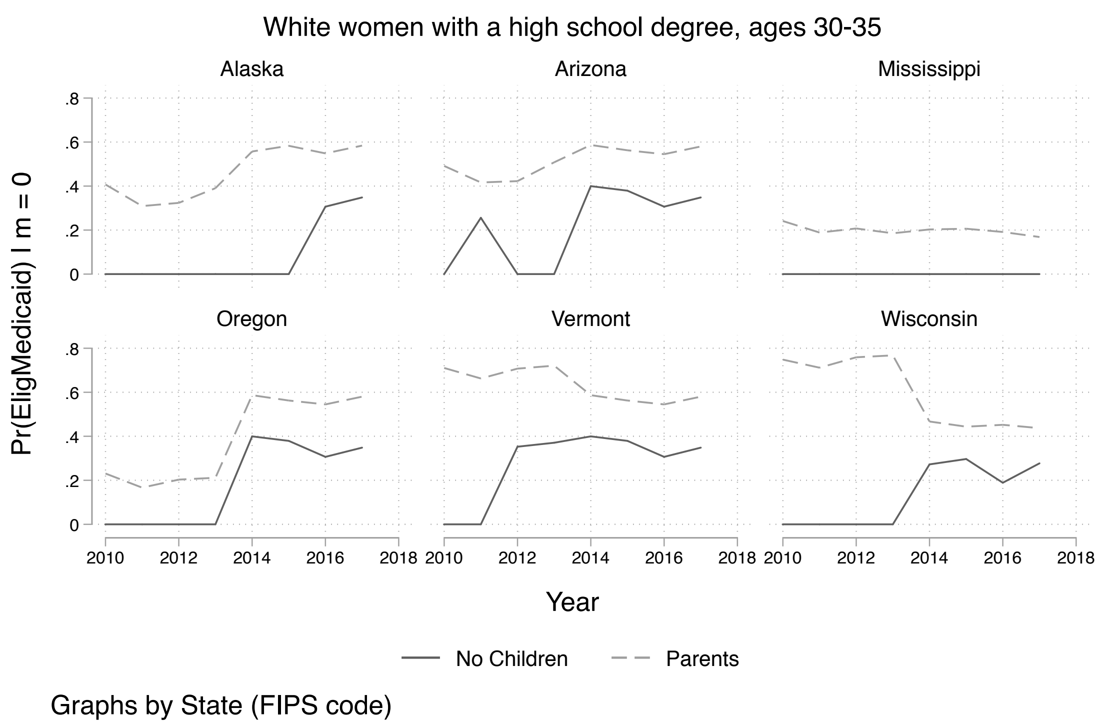
7 Figures

Figure 1: Health insurance coverage by marital status



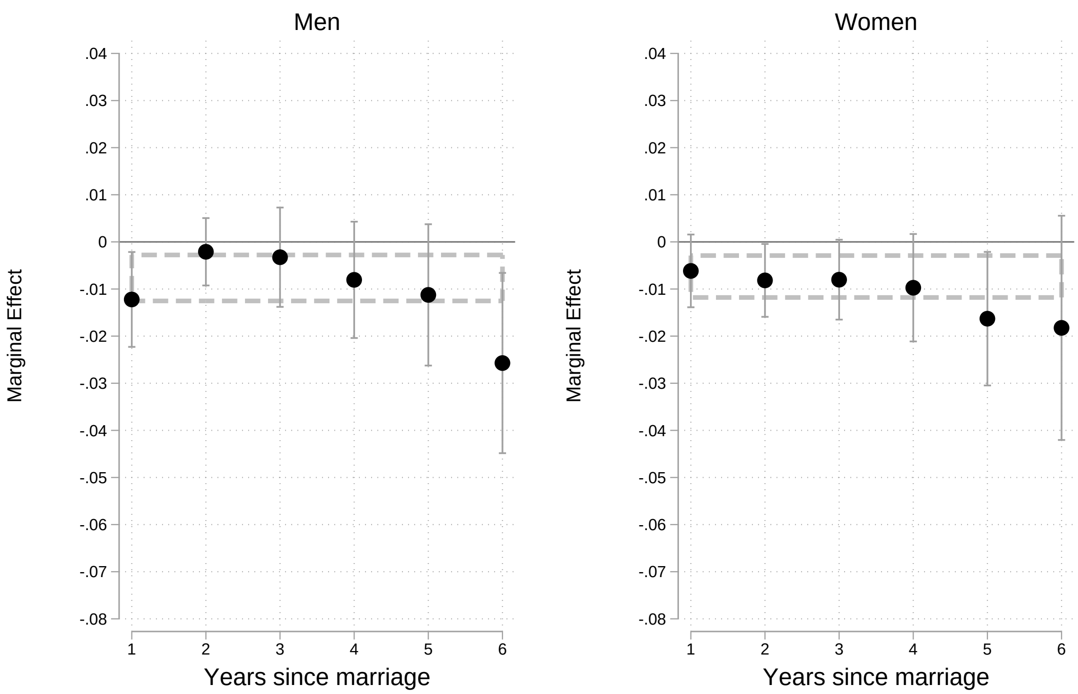
This figure depicts the fraction of men and women that reported having any source of health insurance coverage, by marital status and year, using data from the American Community Survey for adults between the ages of 18 and 64.

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



The figure depicts $PrMedicaid_t$ for a particular demographic group across an array of states in order to illustrate the identifying variation used. $PrMedicaid_t$ 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.

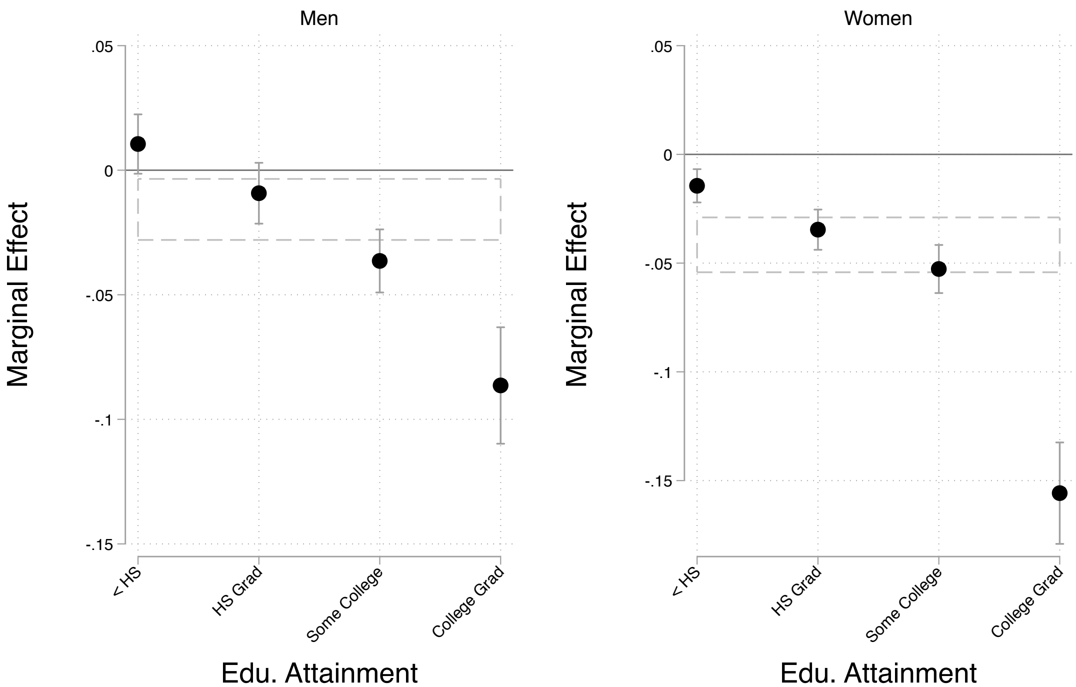
Figure 3: The effect of the probability of Medicaid eligibility at the time of marriage on the rate of divorce, by years since marriage



The estimated coefficients are obtained by weighted least squares regresion, 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 2010-2017. The independent variables are all defined at the time of marriage, such that an individual observation is given by the demographic group and a year of marriage between 2010 and 2017. Each marker comes from a separate regression where observations are pooled by the years since marriage. 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.

Figure 4: The effect of the probability of Medicaid eligibility on marriage rates, by education

(a) Estimated marginal effects



(b) Effect size scaled by mean increase in probability of Medicaid eligibility for states expanding in 2014

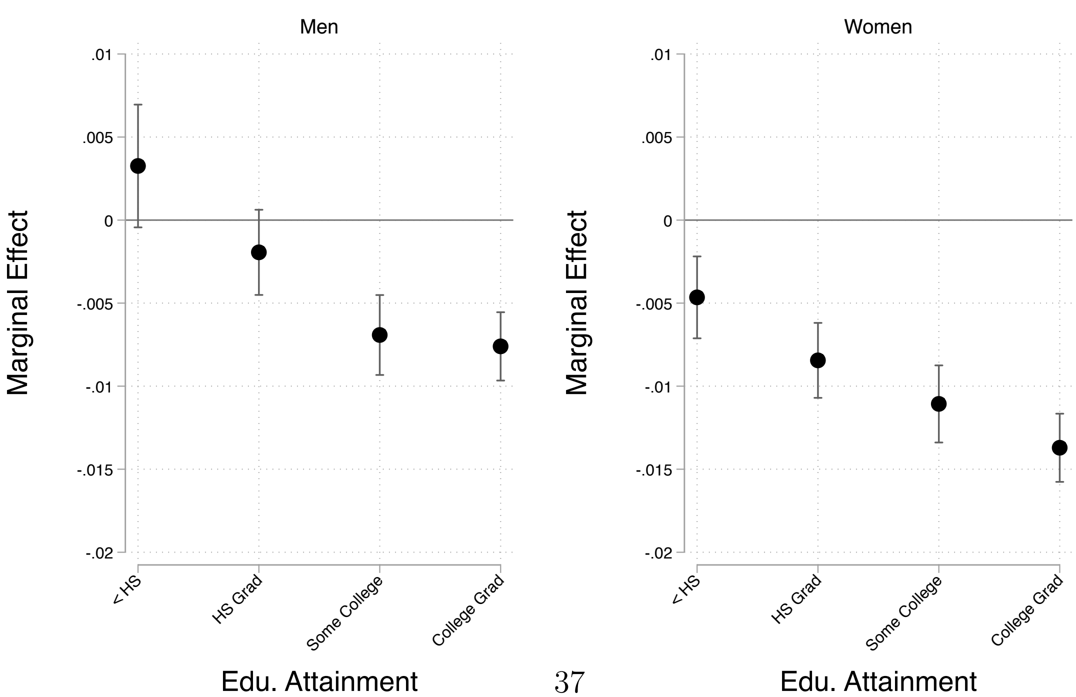
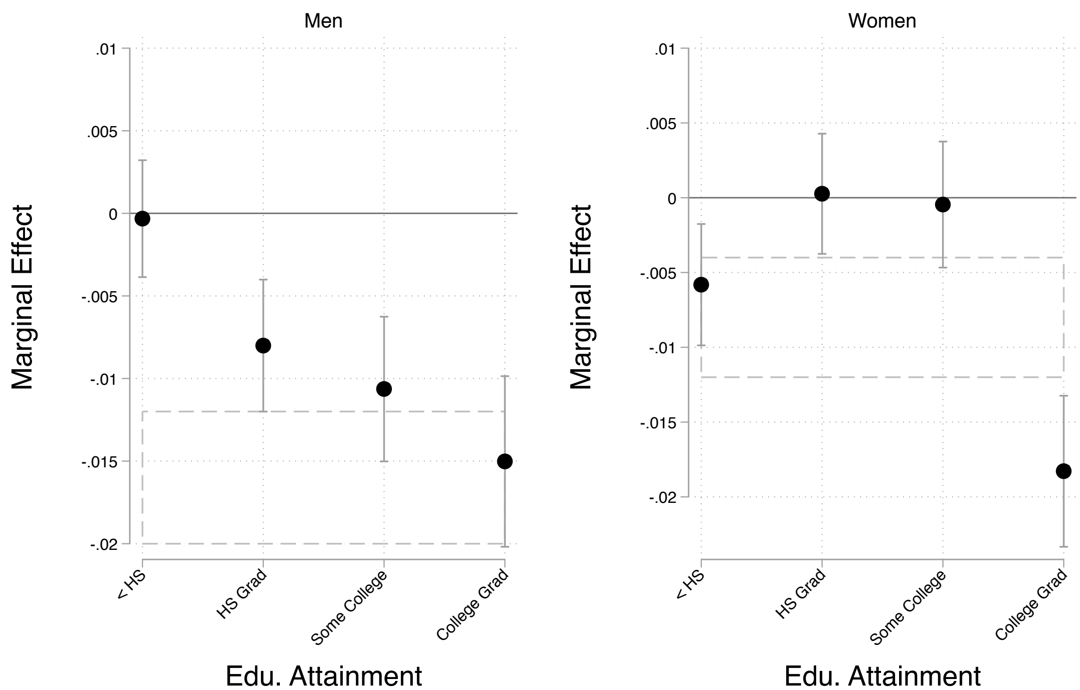
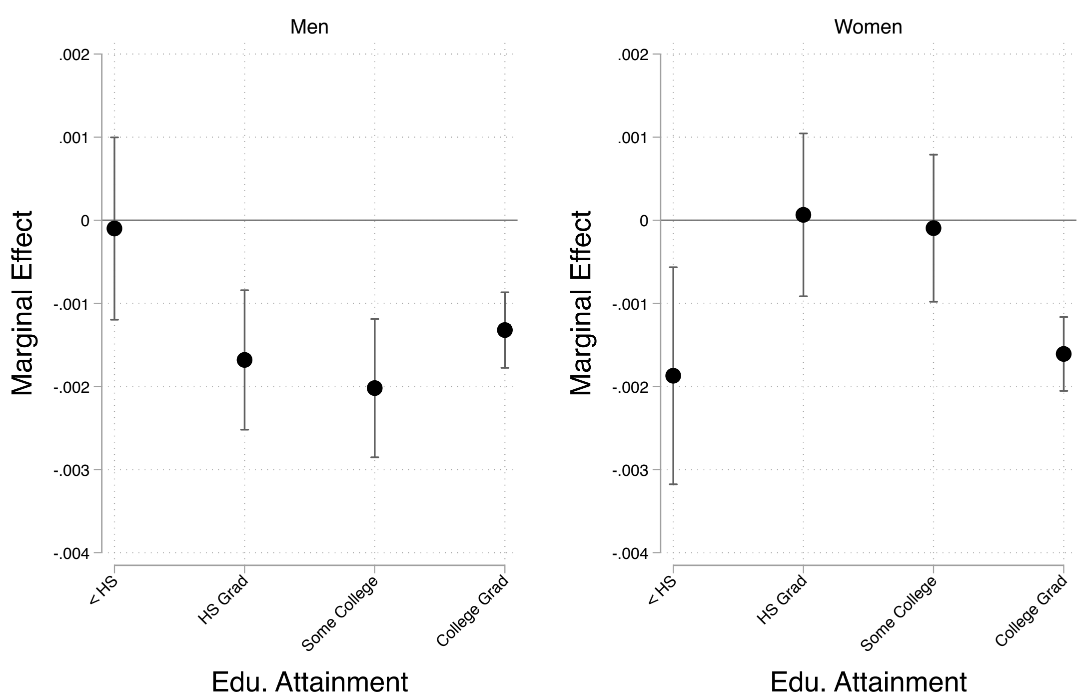


Figure 5: The effect of the probability of Medicaid eligibility on divorce rates, by education

(a) Estimated marginal effects



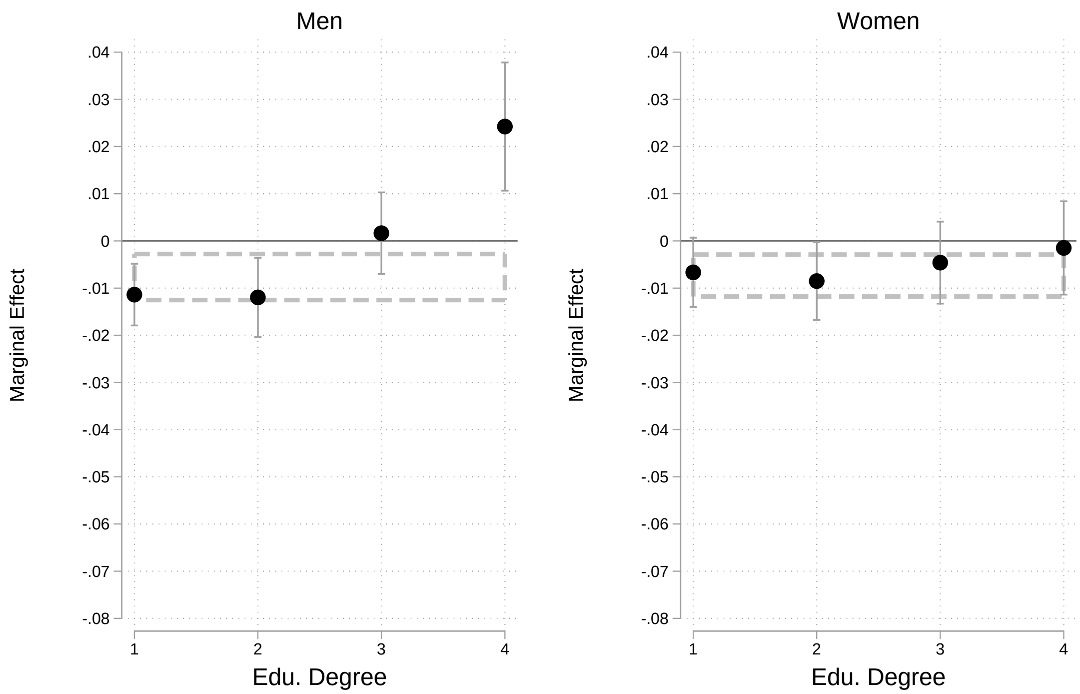
(b) Effect size scaled by mean increase in probability of Medicaid eligibility for states expanding in 2014



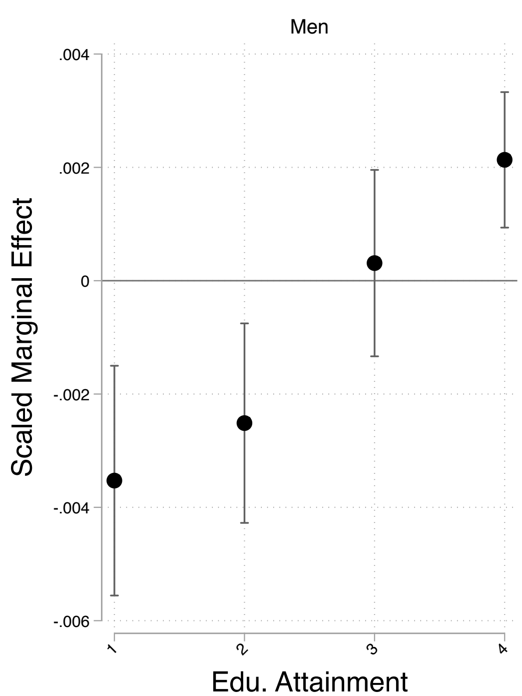
The graph in panel (a) show the total marginal effect of a 100 percentage point increase in the probability

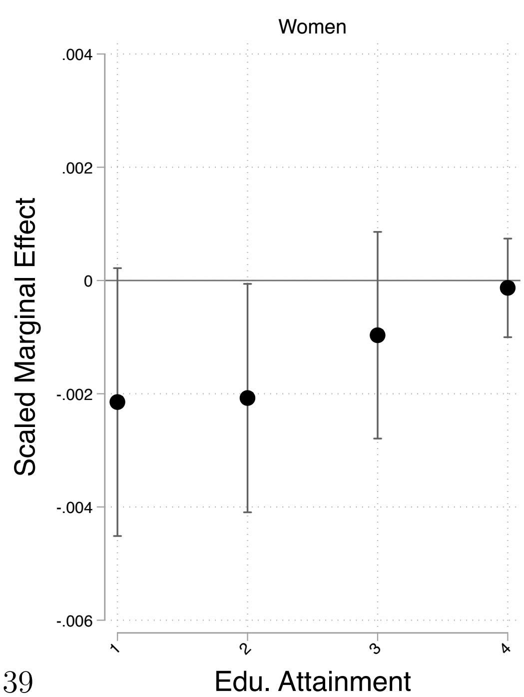
Figure 6: The effect of the probability of Medicaid eligibility in year of marriage on divorce rates, by education

(a) Estimated marginal effects



(b) Effect size scaled by mean increase in probability of Medicaid eligibility for states expanding in 2014





8 Tables

Table 1: Summary Statistics for Full Sample

(a) Demographic Characteristics

	(1) Male mean	sd	(2) Female mean	sd
Age	41.8	13.7	42.2	13.7
White	0.77	0.42	0.76	0.43
Black	0.099	0.30	0.11	0.31
Other aces, grouped	0.13	0.33	0.13	0.34
More than high school	0.51	0.50	0.58	0.49
Parent	0.37	0.48	0.45	0.50
Unemployed	0.059	0.24	0.049	0.22
Not in labor force	0.21	0.41	0.29	0.46
Annual hours worked	1568.4	1047.1	1215.5	983.1
Observations	7439797		7698543	

(b) Insurance Coverage and Medicaid Eligibility

	(1) Male mean	sd	(2) Female mean	sd
Any HI Coverage	0.83	0.38	0.87	0.34
Medicaid Coverage	0.11	0.32	0.14	0.35
Pr(Medicaid)_t	0.10	0.13	0.16	0.19
Observations	7439797		7698543	

(c) Marital Status and Transitions

	(1) Male mean	sd	(2) Female mean	sd
Married	0.50	0.50	0.52	0.50
Divorced	0.100	0.30	0.12	0.33
Marriage Rate	0.033	0.18	0.029	0.17
Divorce Rate	0.015	0.12	0.015	0.12

This table presents descriptive statistics for men and women between the ages of 18 and 64 from the American Community Survey, waves 2010-2017.

Table 2: Duration of Marriage Sample

	(1) Male mean	sd	(2) Female mean	sd
Age_t*	33.9	10.5	32.2	10.4
More than high school	0.57	0.49	0.66	0.47
Parent_t*	0.31	0.46	0.34	0.48
Married	0.84	0.36	0.85	0.36
Pr(Medicaid)_t*	0.075	0.13	0.12	0.19
Year Married (t^*)	2012.9	1.64	2012.9	1.64
Year since t^*	2.28	1.64	2.30	1.64
Observations	470709		478055	

This table presents descriptive statistics for men and women between the ages of 18 and 64 who were married at some point during 2010 and 2017, from the American Community Survey, waves 2010-2017.

Table 3: The effect of the probability of Medicaid eligibility on contemporaneous marriage rates

Marr. Rate in t	Men		Women	
$Pr(Medicaid)_t$	-0.018	-0.022	-0.042	-0.048
	(0.004)	(0.006)	(0.004)	(0.005)
Observations	72,900	72,900	73,433	73,433
Demog. X Year	NO	YES	NO	YES

Robust 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 from the American Community Survey, waves 2010-2017. The dependent variable is the marriage rate for the demographic group, namely the fraction of individuals reporting marriage in the last year relative to the total number of individuals that were not 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.

Table 4: The effect of the probability of Medicaid eligibility at the time of marriage on the rate of divorce

Div. Rate in $t * +i$	Men		Women	
$Pr(Medicaid)_{t^*}$	-0.008***	-0.006**	-0.007***	-0.008***
	(0.002)	(0.003)	(0.002)	(0.002)
Observations	117,784	117,784	114,812	114,812
Demog. X Year	NO	YES	NO	YES

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

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 2010-2017. 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 independent variables are all defined at the time of marriage, such that an individual observation is given by the demographic group and a year of marriage between 2010 and 2017. The controls in Cols. 1 and 3 include dummies for years since marriage, state, educational attainment (four levels), five-year age group at the time of marriage, race (white, black, and other races combined as a third group) and inferred parental status at the time of marriage. Cols. 2 and 4 additionally allow for interactions between the full set of demographic characteristics and years since marriage. Standard errors are clustered at the level of inferred parental status at the time of marriage and state.

Table 5: The effect of the probability of Medicaid eligibility on contemporaneous divorce rates

Div. Rate in t	Men		Women	
$Pr(Medicaid)_t$	-0.011	-0.011	-0.004	-0.005
	(0.002)	(0.002)	(0.002)	(0.002)
Observations	72,900	72,900	73,433	73,433
Demog. X Year	NO	YES	NO	YES

Robust 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 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.

Table 6: Annual Hours of Work

Annual Hours in t	Men			Women		
$Pr(Medicaid)_t$	-12.7	-5.6	4.4	-114.1***	-157.9***	-150.0***
	(16.4)	(13.5)	(12.8)	(15.5)	(15.0)	(14.6)
$Pr(Medicaid)_t^{Sp}$		-9.1	10.5		78.7***	82.3***
		(12.7)	(13.3)		(19.9)	(19.2)
$Pr(Medicaid)_t^{Coup}$			-59.8***			-24.0
			(9.6)			(15.1)
Observations	521,789	521,789	521,789	475,594	475,594	475,594

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

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 outcome variable is annual hours, calculated by multiplying weeks worked in the past year by usual hours per week. Those not working whether due to unemployment or non-participation are assigned zero hours. The controls in each specification include include dummies for year and state, and for each spouse also their educational attainment (four levels), five-year age groups, race (white, black, and other races combined as a third group) and parental status. Standard errors are clustered at the level of parental status and state.

Table 7: Migration inflows and the probability of Medicaid eligibility

Migration Inflow Rate	All		Married	
	Men	Women	Men	Women
$Pr(Medicaid)_t$	0.007	-0.016	-0.009	-0.023
	(0.004)	(0.004)	(0.003)	(0.004)
Observations	84,334	85,790	73,833	75,001

Robust 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 and state, using adults ages 18 to 64 from the American Community Survey, waves 2010-2017. The dependent variable is the fraction of individuals in the demographic group and state that moved from another state in the past year. Cols. 1 and 2 include all individuals regardless of marital status, while Cols. 3 and 4 include only married individuals. The controls in each specification include include dummies for year and state, educational attainment (four levels), five-year age groups, race (white, black, and other races combined as a third group) and parental status. Each of the demographic characteristics are interacted with a full set of year dummies to allow for flexible time trends. Standard errors are clustered at the level of parental status and state.

Table 8: Marital outcomes, migration inflows, and the probability of Medicaid eligibility

	Marr. Rate		Div. Rate		Div. Rate	
	Men	Women	Men	Women	Men	Women
$MigIn_t$	0.179	0.283	0.001	0.000	-0.011	-0.009
	(0.028)	(0.033)	(0.003)	(0.002)	(0.006)	(0.004)
$Pr(Medicaid)_t$	-0.029	-0.058			-0.015	-0.008
	(0.008)	(0.007)			(0.002)	(0.002)
$Pr(Medicaid)_t^*$			-0.006	-0.008		
			(0.003)	(0.002)		
Observations	73,833	75,001	117,784	114,812	73,833	75,001

Robust 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 and state, using adults ages 18 to 64 from the American Community Survey, waves 2010-2017. The dependent variables include marriage rates (Cols. 1 and 2), and divorce rates (Cols. 3-6). Cols. 3 and 4 differ from Cols. 5 and 6 in that the former estimate the effect of eligibility at the time of marriage on divorce rates in subsequent years, and hence all control variables including the rate of in-migration are defined at the time of marriage, whereas the latter estimate the effect of current eligibility on contemporaneous divorce for already-married couples. The controls in Cols. 1-2 and 5-6 include include dummies for year and state, educational attainment (four levels), five-year age groups, race (white, black, and other races combined as a third group) and parental status. Each of the demographic characteristics are interacted with a full set of year dummies to allow for flexible time trends. In Cols. 2-3, there are dummies for the year of marriage and years since marriage, and the demographic characteristics are instead interacted with the set of dummies for years since marriage. Standard errors are clustered at the level of parental status and state.

Table 9: Local economic and political conditions and probability of Medicaid

(a) Marriage rates

Marr. Rate	Men				Women			
$Pr(Medicaid)_t$	-0.022	-0.021	-0.020	-0.020	-0.046	-0.046	-0.044	-0.043
	(0.006)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
$Unemp_t$	0.029	-0.003		-0.003	-0.075	-0.075		-0.077
	(0.016)	(0.014)		(0.015)	(0.020)	(0.020)		(0.021)
$Earnings_t$		-0.028		-0.028		0.000		-0.000
		(0.003)		(0.003)		(0.003)		(0.003)
$DemGov_t$			0.000	0.000			0.000	0.000
			(0.001)	(0.001)			(0.001)	(0.001)
$RepGov_t$			-0.000	-0.001			0.001	0.001
			(0.001)	(0.001)			(0.001)	(0.001)
Observations	72,900	72,900	61,894	61,894	73,433	73,433	62,483	62,483

Robust standard errors in parentheses

(b) Divorce rates and contemporaneous probability of Medicaid eligibility

Divorce Rate	Men				Women			
$Pr(Medicaid)_{t^*}$	-0.014	-0.014	-0.011	-0.011	-0.012	-0.012	-0.011	-0.011
	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)
$Unemp_{t^*}$	0.006	0.004		0.004	0.004	0.004		0.004
	(0.004)	(0.004)		(0.004)	(0.004)	(0.004)		(0.004)
$Earnings_{t^*}$		-0.002		-0.002		-0.000		0.000
		(0.001)		(0.001)		(0.001)		(0.001)
$DemGov_t$			-0.001	-0.001			-0.001	-0.001
			(0.001)	(0.001)			(0.001)	(0.001)
$RepGov_t$			0.002	0.001			0.001	0.001
			(0.001)	(0.001)			(0.001)	(0.001)
Observations	114,398	114,398	109,571	109,571	109,408	109,408	105,008	105,008

Robust standard errors in parentheses

(c) Divorce rates and probability of Medicaid eligbility at time of marriage

Div. Rate	Men				Women			
$Pr(Medicaid)_t$	-0.011	-0.011	-0.011	-0.011	-0.006	-0.006	-0.006	-0.006
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)
$Unemp_t$	0.053	0.042		0.044	0.039	0.037		0.038
	(0.005)	(0.005)		(0.005)	(0.003)	(0.003)		(0.004)
$Earnings_t$		-0.009		-0.009		-0.001		-0.001
		(0.001)		(0.001)		(0.001)		(0.001)
$DemGov_t$			-0.000	-0.000			-0.000	-0.000
			(0.000)	(0.000)			(0.000)	(0.000)
$RepGov_t$			0.000	0.000			0.000	-0.000
			(0.001)	(0.001)			(0.000)	(0.000)
Observations	72,900	72,900	61,894	61,894	73,433	73,433	62,483	62,483

Robust standard errors in parentheses

Panels (a), (b), and (c) are consistent with the regressions estimated in Tables **??** and 5 respectively, using the specification that allows for flexible time trends. Cols. 1-2 and Cols. 5-6 include additional controls for the mean unemployment rate and the mean earnings of the demographic group in a given state and year. Cols. 3-4 and Cols. 7-8 restrict the sample to states that officially participated in the ACA Medicaid expansion at some point during 2010-2017.

A 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.