Economic and social impacts of the Affordable Care Act (ACA)

Dissertation

Presented in Partial Fulfillment of the Requirements for the Degree Doctor of Philosophy in the Graduate School of The Ohio State University

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2020

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Abstract

This dissertation investigates the economic and social impacts of the Affordable Care Act (ACA). The three chapters explore the main theme in the contexts of homeownership, marriage, and as a social safety net in the face of the Covid-19 pandemic.

The first chapter examines the impact of the ACA Medicaid expansion on homeownership among low-income Americans using state, time and income variation in the policy. Medicaid provides direct financial protection in the case of illness or injury, improving financial health and thus lowering barriers to homeownership. Using data from the American Community Survey (ACS), I find that the ACA Medicaid expansion increased the probability of being a homeowner among low-income Americans aged 43-64 by 4 to 8 percent. The marginal household induced into homeownership by the expansion did so with a mortgage. I also explore whether the individual-level behavior changes affect market-level outcomes. Using data from Zillow Research and the Building Permit Survey, I find suggestive evidence that the demand for housing and housing prices increased in areas where large shares of the population became eligible for Medicaid through the ACA, while the supply of new housing units was not affected. My findings suggest that health insurance policy may have important spillovers to other household financial decisions, and that it can impact market-level outcomes.

The second chapter evaluates the extent to which the structure of ACA eligibility and generosity criteria impose marriage penalties or subsidies, and whether the incentives embedded in the ACA affect marriage decisions. The debate on whether the ACA may deter marriage was

heated in popular press. Similar to the U.S. tax income system and transfer programs, the ACA can generate negative or positive economic benefits that affect the decision to marry. These economic benefits are usually called marriage penalties or marriage subsidies. In this study, the ACA marriage penalty or subsidy is defined as the difference between the total premium payment that two individuals would face if they are married, and the total payment if they are unmarried. I exploit the variation in state, time, number of children, and age to investigate the impact of the ACA marriage penalty (or subsidy) on marital status. Using data from the American Community Survey (ACS), I find that the probability of being married drops by 6 percentage points (or 9 percent) with a $1,000 increase in marriage penalty. I also find the ACA marriage penalties have a larger impact on less educated and racial/ethnic groups other than White.

The third chapter studies how access to government subsidized health insurance coverage varies across Black and White people who lost their employer-sponsored insurance (ESI) due to a Covid-19 layoff. I explore the question from three angles. First, I analyze how Covid-19 layoffs interact with gaps in the ACA coverage net to narrow or widen the Black-White disparity in access to health insurance. Second, I evaluate the extent to which Covid-19 federal aid affected the Black-White disparity in access to health insurance. Third, given the uncertainties in the future labor market and continuity of government unemployment compensation, I estimate how the Black-White disparity in access to health insurance will be affected after all unemployment benefits and federal compensation is exhausted in 2021. Using data from the American Community Survey (ACS) and the Current Population Survey (CPS) and a simulation-based approach, I show that the federal unemployment compensation (FPUC) helped reduce the potential increase in the Black-White disparity in health insurance coverage caused by the

interaction between Covid-19 layoffs and existing gaps in the ACA. However, the disparity will increase substantially if the economic recovery is slow and unemployed people exhaust all unemployment compensation in 2021. Improving the ACA provisions is essential to enhance racial/ethnic equity in health outcomes.

Dedication

To my parents, grandma, and Beta --- loved ones in my life

Acknowledgments

I would like to express my sincere appreciation to my advisor, Dr. Lauren Jones, who guided me to the research area in public policy analysis. I learned tremendously from her training, not only about rigorous research methodologies but also the attitude of being a good researcher. I would like to express my heartfelt gratitude to Dr. Tansel Yilmazer, who continually helped me throughout the program in many ways. Her insights and suggestions are so invaluable and always stimulate my analytical thinking. I also deeply thank Dr. Sherman Hanna for all his help throughout the program and especially for co-authoring with me to publish the first journal article in my life. I would also particularly like to thank Dr. Catherine Montalto for her great support through the job market season. Finally, I would like to thank my dear parents and grandma for their unconditional love, and my cat Beta and my friends for their accompany and encouragement in the past four years.

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Fields of Study

Major Field: Consumer Sciences

Minor Field: Economics

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Chapter 1. The ACA Medicaid Expansion and Homeownership

1.1 Introduction

The passage of the Affordable Care Act (ACA) represents the largest evolution in American health care policy since Medicare (Blumenthal et al., 2015). The reform aimed at reducing the uninsured rate by enabling lower income Americans to obtain health insurance coverage. The ACA became effective in 2010, and major provisions became effective in January 2014. As of 2016, about 20 million uninsured nonelderly Americans gained health insurance (Kaiser Family Foundation, 2016). Medicaid expansion – which extends free public health insurance to individuals with income below 138 percent of the Federal Poverty Level (FPL), was a key provision of the ACA – accounted for 60 percent of the coverage gains (Frean et al., 2017).

Millions of uninsured individuals in the U.S. use health care services every year (Coughlin, T.A. et al., 2014). Uninsured patients are charged higher rates by hospitals than insured patients (Xu et al., 2017). Even though uninsured individuals can get some implicit subsidy from private grant programs or Medicaid DSH payments, which are required by federal law to help uninsured individuals with medical care, they still pay one-fourth of their care out-of-pocket (Coughlin, T.A. et al., 2014). Facing a higher burden of health care costs, uninsured individuals are more likely than the insured to sacrifice spending on food and housing, or to encounter other financial problems as a result of medical expenditures[1]. Therefore, besides

[1] https://www.nytimes.com/2016/01/06/upshot/lost-jobs-houses-savings-even-insured-often-face-crushing-medical-debt.html?_r=0

increasing health insurance coverage, the ACA may also have had important financial implications for individuals who gained health insurance. It has been found by recent studies that the ACA Medicaid expansion provided direct financial protection from loss associated with medical expenses and reduced medical collection among individual who gained insurance. As a result, individuals enjoyed improved credit scores and increased access to credit markets (Caswell & Waidmann, 2017; Dobkin et al., 2018; Brevoort et al., 2017; Miller et al., 2018; Hu et al., 2018). Improved financial health can potentially change low-income individuals' financial decisions over a variety of spending categories. In this study, we investigate the effect of the ACA Medicaid expansion on financial decision-making in one important spending category: homeownership.

Owning a home is an important American dream for many people. Homeownership is a major saving mechanism and can protect families from risks of economic adversity, strengthen communities, foster civic pride and provide children with a stable living environment (Herbert et al., 2005). It also represents the largest investment in many people's lives. However, many low-income Americans face barriers to homeownership. Between 2014 and 2019, homeownership rate was about 78 percent among households with median income and above, while it was less than 50 percent among households with less than median income (U.S. Census Bureau, 2019). Low-income individuals face both supply and demand side constraints that prevent them from purchasing a home. On the supply side, the supply of mortgage credit is limited. Lenders often impose credit constraints because of the fear of payment default risk, as low-income individuals are more likely to have worse credit scores and have unstable resources to commit to a stream of mortgage payments. On the demand side, individuals who are more sensitive to financial risk are less likely to want to own a home. Housing is a large durable asset and the transaction costs of

selling are high if moving is necessary (Herbert et al., 2005). These risks may weigh especially heavily on individuals with other financial risks, such as medical expenditure risk.

The ACA Medicaid expansion could potentially affect low-income homeownership through multiple channels. First of all, Medicaid helps reduce medical expenditure risks. With lower likelihood of being financially overwhelmed when getting sick or injured, low-income individuals who gain insurance may feel more confident about becoming homeowners and committing to a stream of mortgage payments. Reduced medical expenditure risk could also change investment portfolio and increase other risky asset holding such as home equity. Second, Medicaid helps reduce out-of-pocket medical expenditures for uninsured individuals with medical expenses. This could increase the resources that can be allocated to non-medical consumption, such as housing. Third, Medicaid can provide direct financial protection with losses associated with medical expenses. With lower probability of being delinquent and have medical debts or collection, low-income individuals who gain health insurance can enjoy improved financial health and better credit scores. With increased access to credit markets and lower barrier of obtaining mortgage credits, low-income individuals might be more likely to enter homeownership.

This study aims to empirically test the effect of the ACA Medicaid expansion on homeownership among low-income Americans. Previous studies about the homeownership mainly focus on factors that are directly related to homeownership decisions, such as housing policies, financial returns to homeownership, and life-cycle stage. Few studies explore the spillover effects of other social policies on homeownership. We contribute to the homeownership literature by examining the effect of the recent largest health insurance expansion on homeownership. In addition, the reduced medical expenditure risks and potential increased risky

asset holding in home equity shed lights on how families balance risk across a variety of domains. We also contribute to the literature on the impact of the ACA Medicaid expansion on individual wellbeing outside of health care use.

We use a difference-in-difference-in-differences (DDD) research design and exploit the state, time, and income variation in Medicaid eligibility after the ACA policy became effective in 2014. Our analysis compares the homeownership status for individuals living in a treatment state who became newly eligible for Medicaid after the expansion, to the homeownership status among three control groups: (i) similar individuals living in an expansion state before 2014, (ii) all individuals living in non-expansion states, and (iii) individuals living in an expansion state who were previously eligible for Medicaid before 2014. These three control groups were not affected by the ACA Medicaid expansion, and thus allow us to control for confounding factors that vary by state, time, and income. Therefore, our estimated effects capture the causal effect of the policy rather than other confounding factors, such as other state laws that became effective at a similar time to the ACA, or time-variant economic factors like the recovery in the housing market.

We particularly explore the heterogeneous effects on different age groups. Compared with younger individuals, middle-age individuals are more likely to experience illness and thus tend to have a higher medical expenditure risk and larger out-of-pocket medical expenses. According to the Medical Expenditure Panel Survey (MEPS), in 2016, the mean medical expenditure per person was $2,985 for individuals aged between 18-44, and $6,406 for individuals aged between 45-64[2]. Therefore, the ACA Medicaid expansion may have larger

[2] Agency for Healthcare Research and Quality. Mean expenditure per person by age groups, United States, 1996-2016. Medical Expenditure Panel Survey.

impact on homeownership among older individuals. Besides, the incentive or capability of buying a home may be different among different age groups, in the aspects of saving for the down payment, life attitude, and bequest motive.

Using data from the American Community Survey (ACS) 2009 to 2017, we find that the ACA Medicaid expansion increased the probability of being homeowners for middle-age individuals (age 43-64), especially for those are older. Being newly eligible for Medicaid increased the probability of being a homeowner by 2 to 4 percentage points. Compared to the baseline homeownership rate, the estimates suggest the policy increased homeownership by about 4 to 8 percent. We also find that the probability of having a mortgage increased by a similar magnitude, while the probability of owing a house without an outstanding loan was unaffected. Hence, the marginal household induced into homeownership by the expansion did so with a mortgage.

If the ACA Medicaid expansion increased the demand for houses among low-income individuals, we would expect to find some changes in the housing market. To test this hypothesis, I use housing data from Zillow Research and Building permit Survey to examine changes in housing price, sales volume and new housing units in areas where a large population was affected by the ACA Medicaid expansion. Using the 2009-2017 zip code or county level housing data, we exploit a similar DDD research design. We compare housing outcomes in expansion state areas that have large shares of uninsured, low-income people, to housing outcomes among three control groups: (i) similar areas in treatment states before 2014, (ii) all areas in control states, and (iii) areas (in expansion states) with low shares of uninsured low-income people. Our results show suggestive evidence that the ACA Medicaid expansion

increased sales volumes and housing prices in those areas where a large population was affected by the expansions. The ACA policy had little impact on the number of new housing units.

The paper is organized as follows: Section 2 addresses the policy details of the ACA Medicaid expansion and the existing literature on homeownership of low-income households. Section 3 describes the data. Section 4 explains the research method. Section 5 shows the estimation results. Section 6 gives a conclusion.

1.2 Background

The Affordable Care Act (ACA) was passed in 2010, and major provisions became effective in January 2014. There were three key provisions of the policy: Medicaid expansion, premium subsidies, and the individual mandate. Medicaid expansion provides free public health insurance to individuals with family income below 138 percent of FPL. Premium subsidies are refundable tax credits available to individuals with income between 100-400 percent of FPL who purchase individual private health insurance. The individual mandate, which was repealed in 2017 and eliminated in 2019, required uninsured individuals to pay a tax penalty. As of 2016, about 20 million uninsured nonelderly Americans gained health insurance. According to Frean et al. (2017), Medicaid expansion produced 60 percent of the coverage gains, premium subsidies accounted for the other 40%, while the individual mandate had little impact.

Before the ACA, Medicaid was mostly available to low-income children and parents, pregnant women, and disabled individuals. Childless adults were generally ineligible for Medicaid in most states. Eligibility was based on family income, and the income threshold varied across states. For example, in 2013 the income eligibility threshold was only 16 percent of FPL in Arkansas, while it was 71 percent of FPL in Washington. Under the ACA, the Medicaid

expansion extended eligibility to all individuals with family income below 138 percent of FPL, regardless of whether they had dependent children or not. However, states were allowed to opt out of the Medicaid expansion. Only 31 states plus Washington D.C. expanded Medicaid by 2017. The remaining states – the non-expansion states – opted not to expand Medicaid[3].

Anecdotal stories have demonstrated the link between health care and housing, which represent the two biggest parts of household consumption. People with higher health care costs may be more likely to sacrifice spending on food and housing. Growth in health care costs have outpaced the economy for decades. The affordability index – a ratio of the average employer-sponsored family health insurance to median family income – rose from below 15 percent to over 30 percent between 1999 and 2016 (Ezekiel et al., 2017). The health care cost burden is more severe among poorer individuals. According to Goldman et al. (2018), 16 percent of the lowest income families (0-138 percent of FPL) has high-burden health spending[4], while less than 1 percent of the higher income families (above 400 percent of FPL) has high burden spending. Over the same period, housing prices and rent also grew faster than average incomes. The median sale price to household income rose from 3.4 to 4.3[5]. The rising health care and housing costs especially hurt low- and moderate- income households who are either struggling or just getting by. While some people postpone or skip medical treatment when they cannot afford it, others may pay for care with the money they need for housing. Gallagher et al. (2019) found that the other ACA provision, premium subsidies, reduced the likelihood of missing rent or mortgage

[3] The states that expanded Medicaid in Jan 2014 are Arizona, Arkansas, California, Connecticut, Colorado, Delaware, Washington DC, Hawaii, Illinois, Iowa, Kentucky, Maryland, Minnesota, Nevada, New Jersey, New Mexico, New York, North Dakota, Ohio, Oregon, Rhode Island, Vermont, West Virginia, Washington. Also, Alaska, Indiana, Louisiana, Michigan, Montana, New Hampshire, Pennsylvania expanded Medicaid from 2014 to 2016. See : https://www.kff.org/health-reform/state-indicator/state-activity-around-expanding-medicaid-under-the-affordable-care-act/

[4] "High-burden" as health spending exceeding 19.5 percentof family income.

[5] https://www.jchs.harvard.edu/blog/price-to-income-ratios-are-nearing-historic-highs/

payments by 25 percent through lowering exposure to out-of-pocket medical expenditure risk. Not being delinquent on home payments will save people from costly eviction or foreclosure. However, they did not specifically examine the effect of the ACA on foreclosure or loss of homeownership.

The ACA Medicaid expansion could influence homeownership through multiple channels. First, through extending health insurance to more low-income individuals, the ACA Medicaid expansion lowers their medical expenditure risk. Social insurance programs like the Medicaid can smooth consumption when an economic shock is idiosyncratic to the household, such as a health problem (Kniesner and Ziliak, 2002). With lower probability of being financially overwhelmed when getting sick or injured, low-income individuals might feel more confident about committing to a stream of mortgage payments and being a homeowner. In addition, the reduced medical expenditure risk could change household investment portfolio, increasing risky asset holding. According to Christian & John (1996), risk-averse individuals will became more risk-averse when an unfair background risk[6] is added to wealth. This leads to a reduced demand for risky assets. Medical expenditure is an increasingly important contributor to financial risk. When this background risk rises, individuals will reduce their exposure to other risks (Goldman & Maestas, 2012). If the opposite is true, then reducing the medical expenditure risk will increase the demand for risky assets. Compared with other risky investments such as stocks and bonds, home equity investment is more attractive because the internal rate of return to homeownership is more favorable (Goodman & Mayer, 2018). Homeownership can help people accumulate assets in many ways, such as tax advantages, borrowing against home equity for greater financial flexibility, encouraging saving, etc. Together, the Medicaid expansion that

[6] Background risk is the risk that cannot be diversified.

lowers medical expenditure risk could potentially encourage low-income individual to obtain homeownership.

Second, the Medicaid can lower out-of-pocket medical expenditure and expand budget constraint. For low-income uninsured individuals with high medical expenditures, gaining health insurance reduces their out-of-pocket medical costs and thus allows more resources to be allocated to non-medical consumption, such as housing (Gruber & Yelowitz, 1999; Levy et al., 2019). One recent study found the other ACA provision, premium subsidies, reduced the likelihood of missing rent or mortgage payments by 25 percent through lowering out-of-pocket medical expenditure (Gallagher et al., 2019). Not being delinquent on home payments will save people from eviction or foreclosure. Hence, the Medicaid expansion could potentially save people from costly default and maintain homeownership.

Finally, the ACA Medicaid expansion could encourage homeownership through increasing the likelihood of accessing mortgage credits among low-income individuals. Low-income individuals are more likely to face difficulty with paying medical expenses and have payment deliquency, hurting their credit scores and limiting their access to credit markets. Recent studies have shown that the ACA Medicaid expansion, which provides direct financial protection to more low-income individuals, improved their financial health in many ways. Low-income individuals who gained health insurance were less likely to be delinquent, had medical collection and filed bankruptcy (Baicker et al., 2013; Finkelstein et al., 2012; Hu et al., 2018). Brevoort et al. (2017) found the ACA Medicaid expansion improved individual's credit score. As a result, credit card interest rates, auto loan rates and mortgage rates significantly dropped in treatment states, compared to control states. These findings suggest that individuals who gained insurance should be more willing and able to borrow through credit markets. Miller et al. (2018)

confirmed this: they found that the ACA Medicaid expansion increased credit card borrowing and auto borrowing. However, they did not examine mortgage borrowing.

Compared with younger individuals, middle-age individuals especially those who are older usually have higher medical expenditure risk and larger out-of-pocket medical costs. Thus, the ACA Medicaid expansion might have larger impacts on middle-age individuals. Besides, the incentive or capability of buying a home may be different among different age groups, in the aspects of saving for the down payment, life attitude, and bequest motive. One recent study shows new loan originations are shifting from younger borrowers to older borrowers because older borrowers are more reliable on loan payments (Brown et al., 2019). Together, middle-age individuals might be more likely than younger individuals to obtain homeownership and access the mortgage credits under the ACA Medicaid expansion.

1.3 Data

We use the 2009-2017 American Community Survey (ACS)[7] to examine the effects of the ACA Medicaid expansion on low-income homeownership and mortgage status. The ACS is the largest household survey in the US conducted annually by the Census Bureau. The ACS provides information about demographics, housing, as well as social and economic characteristics for respondents living in different geographic areas. Because of its large sample size (approximately 3 million per year), the ACS is the primary source recommended by the Census Bureau for the study of health insurance at the state- and local-levels (Finegold and Gunja, 2014). The ACS is also largely used in housing research because the questionnaire collects data on housing

[7] Ruggles S, Flood S, Goeken R, et al. IPUMS USA: Version 7.0 [dataset]. Minneapolis, MN: Integrated Public Use Microdata Series; 2017.

characteristics, such as ownership, mortgage status, housing cost components, and dwelling characteristics (Wardrip and Pelletiere, 2008).

We restrict the sample to household heads with family income below 138 percent of FPL[8]. In the ACS, housing outcomes are surveyed at the household level, where a "household" includes all members living in the same dwelling. For example, if a house is owned, not rented, a related family member, such as an adult child, or an unrelated member such as roommate or boarder living with the household head will all be labeled as "homeowners". Only keeping household heads, who is the primary person making housing decisions, we avoid overweighting larger households in the case of counting cohabitating non-homeowners. We also limit the sample to non-elderly individuals aged between 27 and 64, who are most likely be affected by the ACA Medicaid expansion[9].

The main dependent variables of interest are measures of homeownership and mortgage status. In the ACS, respondents not living in group quarters were asked: "Is this house, apartment, or mobile home owned by you or someone in this household?" We created a dummy variable that equals one for individuals who indicated that the house was owned. To identify mortgage status, the ACS asks homeowners whether the housing unit was owned without a loan or was encumbered by a mortgage, loan, or other types of debt[10]. We create a dummy variable that equals one if the individual answered that the house was encumbered by mortgage or other debt, and zero otherwise[11].

[8] In Medicaid expansion states, individuals with family income above 138 percent of FPL are eligible for premium subsidies.
[9] Individuals over 65 are eligible for Medicare and young adults up to age 26 can be included in their parents' private health insurance under the ACA.
[10] Mortgages includes all types of loans secured by real estate, including reverse mortgages. Home equity loan is usually been treated as second mortgage. The ACS asks a separate question for second mortgage.
[11] Having mortgage equals zero for renters.

We also use data from Zillow Research and the Building Permit Survey to examine changes in housing price, housing demand and supply in local housing markets after the ACA became effective. Zillow Research collects time-series housing data across different geographic areas, such as state, county, city, zip code, and neighborhood. It has rich housing outcomes, including a home value index (ZHVI), a rent index (ZRI), sales listing, sales volumes, inventory, etc. The main variables used in this study are sales volume and home value index (ZHVI). Sales volume is the number of homes sold during a given year[12]. ZHVI is a smoothed, seasonally adjusted measure of the median estimated home value across a given region and housing type. Different housing outcomes are available at different levels of geography. We obtain Zillow housing outcomes at the zip code level from 2009-2017. The original Zillow data includes 15,530 zip codes for ZHVI and 30,995 zip codes for sales volume. We include the matched 15,530 zip codes in my sample[13]. The Building Permit Survey is conducted by the Census Bureau. The survey provides information on the number of total new housing units authorized across time and different geography. The smallest geographic area available is county. The survey includes about 3,000 counties[14]Building permits are classified as 1-unit building, 2-unit buildings, 3-4 unit buildings, and 5+ unit buildings[15]. We obtain the number of building permits at the county level from 2009-2017. We calculate the total number of new housing units by multiplying the number of building permits and the number of housing units in the corresponding

12 The original data includes the number of homes sold during a given month. I calculate the annual sales volume by summing up all months' sales volume.
13 There are about 42,000 zip codes in the U.S. About 85 percent of the state population living in those 15,530 zip codes.
14 There are about 3,142 counties in the U.S.
15 The exact number of housing units is available for 5+ unit buildings.

building permit. These data can show whether housing supply was also impacted by the ACA Medicaid expansion.

1.4 Methods

We use a difference-in-difference-in-differences (DDD) research design to examine the effect of the ACA Medicaid expansion on homeownership and mortgage status. The ACA Medicaid expansion policy varies by time, geography, and family income. Using the DDD approach, we identify the effects of the policy by comparing individuals who are most likely to be treated by the policy (newly eligible individuals in treatment states) to those who are less likely to be affected.

Under the ACA, states have the right to decide whether to expand Medicaid or not. By the end of 2017, 31 states and the District of Columbia had expanded Medicaid while the others had not. However, we do not simply define the treatment status of each state by its expansion status as of 2014. Among states that expanded Medicaid, some states had fully expanded Medicaid to all individuals with incomes around or higher than 138 percent of the poverty line before 2014. These are Delaware, Massachusetts, New York, Vermont, and Washington D.C. Some states ("early expanders") already expanded Medicaid to both parents and childless adults in a very significant way before 2014. These states are Arizona, Connecticut, Hawaii, and Minnesota. Some states expanded Medicaid after 2014: Michigan in February 2014, New Hampshire in August 2014, Pennsylvania in January 2015), Indiana in February 2015, Alaska in September 2015, and Montana in January 2016. In the analytic sample, we exclude states that

expanded Medicaid after January 2015[16] and "early expanders". Our treatment states include 19 states that expanded Medicaid in 2014. Our control states include 19 states that have not expanded Medicaid by the end of 2017. I also place 5 states that fully expanded Medicaid before 2014 in the control group because they are not affected by the 2014 ACA Medicaid expansion.

Most studies of the ACA Medicaid expansion use difference-in-differences (DD) method to compare individuals in states that did and did not expand Medicaid under the ACA (Sommers et al., 2014; Black and Cohen, 2015; Wherry and Miller, 2016). My study differs from these is that we also consider a third difference that allows me to better target the policy variable: newly eligible status based on family income. The ACA Medicaid expansion mostly affects individuals who became "newly eligible" in expansion states. Individuals who were eligible for Medicaid before the ACA are less affected by the policy change. With a DDD, we can tease the effect of the ACA Medicaid expansion for the population who are most likely treated. In other words, we are not only comparing individuals in states that expanded Medicaid to those did not, but also comparing individuals who are newly eligible for the Medicaid to those who were previously eligible. This eliminates worries that confounding factors at the state level – for example, the contemporaneous housing market recovery, or other state laws that became effective at a similar time – could bias results. In addition, we use data for more time points, five years before the ACA (2009-2013) and four years after (2014-2017). Using multiple time points allows me to examine the common trends assumption in the DD or DDD research design. We estimate the following basic model:

[16] Michigan and New Hampshire are defined as treatment states. Pennsylvania, Indiana, Alaska, and Montana are excluded from the sample.

$$y_{inst} = \beta_0 + \beta_1 TREAT_S * POST_t * NEWLY\ ELIG_n + \beta_2 TREAT_s + \beta_3 POST_t + \beta_4 NEWLY\ ELIG_n + \beta_5 TREAT_s * POST_t + \beta_6 TREAT_s * NEWLY\ ELIG_n + +\beta_7 POST_t * NEWLY\ ELIG_n + X_{inst}'\gamma + e_{inst} \quad (1.1)$$

In equation 1.1, y_{inst} is a dummy variable that equals one if the individual *i* in eligibility group *n* lives in state *s* at year *t*, is a homeowner or has a mortgage. $TREAT_s$ is a dummy equals one if the individual lives in a treatment state and $POST_t$ is a dummy equals one if the observation comes from 2014 or later. $NEWLY\ ELIG_n$ equals 1 if the individual would be newly eligible for Medicaid if her state had chosen to expand Medicaid (regardless of whether it actually did)[17]. The triple interaction term $TREAT_S * POST_t * NEWLY\ ELIG_n$ is the main variable of interest. It equals one if an observation comes from 2014 or later, lives in a treatment state and is newly eligible for Medicaid. The coefficient β_1 is the key coefficient of interest, which represents the DDD treatment effect of the ACA Medicaid expansion[18]. The vector X_{inst} includes the controls: age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, employment status, and housing price index (ZHVI) for the state of residence.

"Common trends" is an important assumption under the DD and DDD research design methods (Angrist and Pischke, 2008). Common trends assume that the trends in the outcome

[17] "Newly eligible" status is defined by the 2013 Medicaid income eligibility threshold in each state. Individuals with income higher than the 2013 eligibility threshold are defined as "newly eligible" for the Medicaid under the ACA.

[18] $TREAT_s$ captures the baseline difference in the outcomes between treatment states and control states. $POST_t$ captures the change of the outcomes that are common to all states. $NEWLY\ ELIG_n$ captures the baseline difference in our outcomes between newly eligible individuals and previously eligible individuals. $TREAT_s * POST_t$ captures the effect of the ACA Medicaid expansion on previously eligible individuals. $TREAT_s * NEWLY\ ELIG_n$ captures the baseline difference in the outcomes between newly eligible individuals living in treatment states and all other individuals. $POST_t * NEWLY\ ELIG_n$ captures the difference in the change of outcomes between individuals who would be newly eligible for Medicaid if living in treatment states and previously eligible individuals that are common to all states.

would be similar among treatment states and control states (and previous eligible group and newly eligible group) in the absence of the policy intervention. The violation of this assumption will cast doubt on the "true" causal effect. For example, if there is an upward trend of the homeownership rate among newly eligible individuals in treatment states before the ACA was effective, a positive and significant DDD estimate is not enough to conclude that the increase in homeownership is caused by the ACA Medicaid expansion. Adding a state specific time trend can ease the strict "common trends" assumption by allowing different trends of the homeownership among states. If the DDD estimate is still positive and significant after controlling for state specific time trend, the causal inference is more reliable. Therefore, in a separate model, I control for state specific time trend. We also estimate a more flexible model:

$$
\begin{aligned}
y_{inst} = {} & \beta_0 + \beta_1 TREAT_S * POST_t * NEWLY\ ELIG_n \\
& + \beta_2 STATE\ FE + \beta_3 YEAR\ FE + \beta_4 NEWLY\ ELIG_n \\
& + \beta_5 STATE\ FE * YEAR\ FE + \beta_6 STATE\ FE * NEWLY\ ELIG_n + +\beta_7 YEAR\ FE * NEWLY\ ELIG_n \\
& + X_{inst}'\gamma + e_{inst} \quad (1.2)
\end{aligned}
$$

The model is similar to the main model, except that we use a vector of state dummies $STATE\ FE$ instead of a simple dummy $EXPANSION_s$, and use a vector of year dummies $YEAR\ FE$ to substitute $POST_t$. Compared to the basic model, which captures the baseline differences in homeownership by states' treatment status and the average changes of homeownership between pre- and post-ACA period, the flexible model captures the baseline differences in homeownership between each specific state and the changes of homeownership by each year. The flexible model also includes a set of state-year fixed effects (the interaction term of state

dummies and year dummies), which control for state-year specific effects on the outcomes. For example, there could be a state-specific event or policy that is unrelated to the ACA Medicaid expansion affected the homeownership and mortgage status in that state. All models are estimated using OLS.[19] Standard errors are clustered at the state level.

We then use the Zillow data at zip code level to examine the changes in housing price (ZHVI) and sales volume in the areas where a large low-income population was affected by the ACA Medicaid expansion. Our strategy is to exploit a similar DDD research design, with the third difference now coming from the share of low-income uninsured people in zip codes.

To measure the share of low-income uninsured people in each zip code, we use the statistics from the ACS 5-year estimates. The ACS provides multiyear estimates of aggregate socioeconomic and housing characteristics for smaller geographic areas such as zip codes. We obtain the ACS 2009-2013 5-year estimates of health insurance status by income to poverty ratio and age at the zip code level. For each area, the data provides the number of individuals between ages 19-64 with income below 138 percent of FPL who are uninsured. We calculate the share of uninsured individuals age between 19-64 with income below 138 percent of FPL for each zip code. Those people are Medicaid's target. We then define the top quartile zip codes, which have the highest share of uninsured, low-income population, as my potential treatment group. We estimate the model 1.3:

$$y_{ahst} = \beta_0 + \beta_1 TREAT_S * POST_t * TOP_h$$

[19] We understand that OLS has some theoretical disadvantages when dependent variable is dichotomous. Alternative methods such as logit or probit have been recommended. However, OLS is commonly used in economics. For estimation of effects, some empirical evidences have shown that there are little differences between two methods (e.g., Hanna and Lindamood, 1985; Noreen, 1988).

$$+ \beta_2 TREAT_s + \beta_3 POST_t + \beta_4 TOP_h$$

$$+ \beta_5 TREAT_s * POST_t + \beta_6 TREAT_s * TOP_h + + \beta_7 POST_t * TOP_h + e_{inst} \quad (1.3)$$

y_{ahst} is the housing market outcome (ZHVI or sales volume) at zip code *a* in treatment group h in state *s* at year *t*. $TREAT_s$ is a dummy equals one if the individual lives in a treatment state and $POST_t$ is a dummy equals one if the observation comes from 2014 or later. TOP_h equals 1 if the zip code had large share of low-income uninsured people between 2009 and 2013. The triple interaction term $TREAT_S * POST_t * TOP_h$ is the main independent variable of interest. It equals one if the zip code had large share of low-income uninsured people is in a treatment state and comes from 2014 or later. The coefficient β_1 is the key coefficient of interest, which represents the DDD treatment effect of the ACA Medicaid expansion.

Finally, we use the Building Permit Survey to examine the changes of housing supply in the areas where a large low-income population was affected by the ACA Medicaid expansion. The number of building permits are available at county level. We define the treatment and control counties using the similar way as we define the treatment and control zip codes above.[20]

For all housing market outcomes, we analyze the basic model, basic model with state specific time trend, and a flexible model.

[20] I use the ACS 5-year estimates at county level. Treatment counties are those with high share of uninsured and low-income population in treatment states.

1.5 Results

We first show the result of the full sample. We then stratify the sample by age because the ACA Medicaid expansion can have heterogeneous effects on different age groups. The sample is divided into five equal size age groups: 27-34, 35-42, 43-50, 51-57, and 58-64.

1.5.1 Descriptive summary

Table 1.1 presents the descriptive statistics (pre-ACA period) for the sample by state's treatment status. The individual characteristics, in terms of age, race/ethnicity, marital status, household size, number of children, education, employment status, and income are similar between the two groups. In treatment states, 69 percent of the sample are newly eligible for Medicaid, while 75 percent of the sample in control states would be newly eligible if those states expanded Medicaid. This is because in states that decided not to expand Medicaid, the pre-ACA income eligibility threshold is usually much lower[21], and thus more people would have become newly eligible if the state had expanded Medicaid to all individuals with income below 138 percent of FPL. There is also a lower percentage of homeowners and higher housing prices in treatment states, compared to control states. This is mainly driven by California (a treatment state). The results were balanced if California was excluded [22].

1.5.2 Estimated effects of Medicaid expansion on insurance coverage

The underlying assumption of this study is that Medicaid expansion increased the access to health insurance among low-income individuals, and thus influenced their homeownership

[21] Table A.1 shows the states' Medicaid income eligibility thresholds between 2009-2013 and adoption status by December 2017.
[22] DDD estimates are not sensitive regarding to the inclusion or exclusion of California.

decisions. Many studies have shown that Medicaid expansion increased the health insurance coverage (Wherry and Miller, 2016, Kaestner et al., 2017; Frean et al., 2017; Hu et al., 2018). We validate the first stage using my DDD method and data for more time points. The results are reported in Table A.2. We find that Medicaid expansion increased the probability of being covered by Medicaid by about 6 to 9 percentage points for the full sample. The estimates are similar to those in other studies. The results for different age groups show that the Medicaid expansion increased the probability of having Medicaid by about 8 to 10 percentage points among the 43-50 age group. Among the 58-64 age group, the estimated effect decreased to about 4 percentage points and becomes insignificantly different from zero when using the flexible model.

1.5.3 Estimated effects of Medicaid expansion on homeownership

Figure 1.1 shows the time trends for homeownership among treatment and control groups for younger individuals and middle-age individuals. As can be seen, the homeownership trends among treatment and control groups are parallel before 2014 when the ACA became effective. From 2014 and later, the homeownership rate increased slightly more rapidly among treatment group (individuals in treatment states who are newly eligible for Medicaid). Table 1.2 presents the estimated DDD effect of the ACA Medicaid expansion on homeownership for the full sample and different age groups. We start with the results of the basic model, which estimates equation (1). There is a null effect of the policy on homeownership among the full sample. However, we observe an increasing effect across age groups. The estimated effect increases from an insignificant -0.3 percentage points among the youngest group to a significant 3.3 percentage

points among the 58-64 age group. On a base homeownership rate of 56 percent among the oldest group, this represents a 5.9 percent increase in the homeownership (3.356 = 0.059).

We then add linear state specific time trends. Pre-existing homeownership trends over time in states may cause me to overestimate the treatment effect. Our results are still robust after controlling for the state specific time trends. The estimated effect increases from an insignificant 0.2 percentage points among the youngest age group to a marginal significant 1.5 percentage points among 43-50 age group, then to a significant 4.3 percentage points among the 58-64 age group. On a base homeownership rate of 42 percent among 43-50 age group and 56 percent among the 58-64 age group, these represent a 3.6 percent and 7.7 percent increase in the homeownership (1.5/42 = 0.034; 4.3/56 = 0.077).

Finally, we estimate a flexible model, which includes state fixed effects, year fixed effects, and state year fixed effects. The estimated results still hold, but are less significant and with smaller magnitude. The estimated effect for the 58-64 age group reduces to 2.5 percentage points, which represents about 4.5 percentage increase in the probability of being homeowners.

We also plot the estimated coefficients by age group, using the preferred model (basic model with control of state specific time trends). Figure 1.2 shows the effect increases monotonically with age.

1.5.4 Estimated effects of Medicaid expansion on mortgage results

If the ACA Medicaid expansion increased homeownership through improving their financial health, such as being more able to access the credit market, we would expect to find the probability of having mortgage increased at the same time when homeownership increased. We conduct similar analyses for the mortgage status and Table 1.3 shows the results. We find very

similar results compared with homeownership. The basic model and model with linear state specific time trends show that the ACA Medicaid expansion increased the probability of having a mortgage by about 2 percentage points among the 43-50 age group to 4.3 percentage points among the 58-64 age group. On a base mortgage rate of 22 percent among the 43-50 age group and 24 percent among the 58-64 age group, these effects represent about 9.1 percent and 17.9 percent increases mortgage rates, respectively. Figure 1.3 plots the estimated coefficients by age group. We again observe a monotonic increasing effect with age.

To further confirm that the increase of homeownership is caused by the increasing ability to get a mortgage because of the ACA Medicaid expansion, we also examine the probability of owning a home outright. In Table 1.4, we did not observe any significant and large effect of the policy on owning home outright.

1.5.5 Housing market results

Figure 1.4 shows the trends of annual sales volume by treatment and control zip codes. Sales volume in treatment zip codes (high uninsured rate zip codes in treatment states) remained flat until 2015, one year after the ACA was effective, while sales volume in control zip codes all started to jump in 2012, when the housing market rebounded. However, the housing price in treatment states was much higher than the control states, as shown in Figure 1.4. Hence, people living in treatment zip codes are those with very low incomes who face higher housing prices. We observe an increase in sales volume in 2015 for this group, which coincides with the ACA Medicaid expansion. Table 1.5 column (1) shows that the estimated policy effect on sales volume was positive but not statistically significant when using basic model and model with state

specific time trends. However, the flexible model shows that treatment zip codes had a significant 26.6 percent increase in sales volume after the ACA Medicaid expansion.

Figure 1.5 shows the trends of the housing price index (ZHVI) by treatment and control zip codes. Housing price increased more rapidly in treatment states than control states, but both zip codes with high and low uninsured rate in treatment states faced similar increasing trends. The estimated results in table 5 column (2) show that housing price increased by about 12-14 percent among the treatment zip codes after the ACA Medicaid expansion.

Figure 1.6 shows the time trends of new housing units authorized by building permits for the treatment and control counties. The number of new housing units increased more rapidly among counties with a large share of the low-income uninsured people, compared to counties with small share of low-income uninsured people. The trends started in 2012 when the housing market rebounded. The number of new housing units increased more after 2014 in treatment counties, compared with control counties. The regression results in table 1.5 column (3), however, show a positive (about 1.8-3.2 percent increase) but insignificant effect of the ACA Medicaid expansion on the number of new housing units.

1.5.6 Falsification check

We first conduct falsification tests by running analysis on two groups of untreated individuals: (1) older individuals (age 70+) and (2) individuals with family income greater than 138percent of FPL. The ACA Medicaid expansion mainly targets at non-elderly adults since individuals 65 and over are all eligible for Medicare. Individuals with family income greater than 138percent of FPL are not eligible for Medicaid. Those with income below 400percent of FPL are eligible for premium subsidies, however, there is no state variation in eligibility. If my identification strategy

reveals the causal effect of the ACA Medicaid expansion on homeownership, we should not expect any significant changes in homeownership or mortgage status among these two ineligible groups.

Table 1.6 shows the DDD estimates of individuals with age above 70. We exclude individuals with age between 65 and 70 because they may be under 65 in 2014 when the ACA became effective. We find some positive and significant estimated effects of Medicaid expansion on homeownership among some models. However, the results are not robust across different models. The estimated effects for having mortgage for this group are not statistically significant and with magnitude close to zero. Hence, individuals over age 70 did not change their access to the mortgage credits after then ACA policy became effective.

Table 1.7 and Table 1.8 show the estimated effect of the ACA Medicaid expansion on homeownership and mortgage status among higher income individuals. We use a difference-in-differences (DD) method because there is no variation in eligibility status. In other words, we compare higher-income individuals in Medicaid expansion states with those in non-expansion states before and after the ACA became effective. The results show significant effects in some models. However, the estimated effects are not robust across models and have very small magnitude. In addition, there is no pattern that the effects are larger among middle-aged individuals than younger individuals.

To confirm that our result is not driven by other factors that might coincide with the treatment status, we run a placebo test by distributing the treatment status (either state treatment status or newly eligible status) randomly across individuals and re-estimate the DDD effects. In other words, some individuals in the control group (not treated) will be re-coded as being treated, and some individuals in the treatment group (actually treated) will be re-coded as not being

treated. We replicate this random assignment 1000 times for each individual, and then run 1000 regressions where we use each of the randomly generated treated status variable in place of the actual treatment status. We do not expect the average placebo estimate to be significantly different from zero, and we expect our actual estimate to be in the tail of the distribution of placebo estimates. Figures 1.7 and Figure 1.8 show the distribution of the 1000 re-estimated DDD effects for individuals aged between 58-64, who were found to have largest homeownership gain after the ACA Medicaid expansion. The t-test results show the estimates are not significantly different from zero, and our estimated effects (4 percentage points) are above the 95percent confidence interval.

1.6 Conclusion

This paper evaluates the effect of the ACA Medicaid expansion on low-income homeownership decisions. Health care and housing are the two biggest components of household consumption. Health care costs and housing costs have both grown faster than the economy or income over decades, causing higher financial burden. Low-income uninsured individuals with higher medical expenditure risk and out-of-pocket medical expenditures may be more likely to have less resources available for housing or have worse credit scores which increase the barrier of accessing credit markets and obtaining a mortgage. The ACA Medicaid expansion allowed millions of low-income Americans to gain health insurance, greatly lowered the medical expenditure risks and provided important financial protection. Recent studies have shown that the ACA Medicaid expansion improved financial health such as fewer medical debts or collection and better credit scores. With lower medical expenditure risks, more resources for

housing, and easier access to credit markets, low-income individuals should be more willing and able to obtain mortgage and obtain homeownership.

Using the ACS data from 2009 to 2017, we examine whether low-income individuals who were newly eligible for Medicaid under the ACA were more likely to be homeowners, compared to individuals who were less likely affected by the ACA. We also examine whether the change of homeownership was associated with the change of accessing a mortgage. We then use housing market data from Zillow research data and the Building Permit Survey over the same period to examine potential changes in housing price, and demand and supply of housing units in the areas where large populations are affected by the ACA Medicaid expansion.

Because individuals with different ages are different in health or medical expenditure risk and life cycle, we particularly investigate the heterogeneous policy effects on homeownership by age. We find that the ACA Medicaid expansion, which became effective in January 2014, increased the probability of being homeowners by about 2 to 4 percentage points, among middle-age individuals aged over 43, with larger effects on individuals aged between 58 and 64. The probability of having a mortgage also increased by a similar magnitude, while there was no change in outright homeownership. Hence, the marginal household induced into homeownership by the expansion did so with a mortgage. This result is consistent with the recent work documenting that the ACA Medicaid expansion increased access to credit markets, reduced interest rates for low-income individuals in treatment states, and increased credit card and auto borrowing.

There are some suggestive evidences that the sales volume and housing price increased in the areas where a large population was affected by the ACA Medicaid expansion. Housing price index (ZHVI) increased by about 12-14 percent in the areas with large population benefitted

from Medicaid expansion. Sales volume increased by 8-26 percent in those areas. However, the results are not robust across models. The estimated policy effect on the number of new housing units supply is very small and is not statistically significant in any model.

This study contributes to the recent literature that connects health care cost and housing decisions. The results of this paper suggest that health insurance policy may have important spillovers to other household financial decisions, such as homeownership, and that it can impact market-level outcomes. Not only housing policies, financial returns, and life-cycle stage affect the homeownership decision, but also other social policies such as health insurance policy can have a significant impact. On the other hand, health insurance policy might also influence many aspects of family wellbeing, other than changing health care utilization and overall health.

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Table 1.1. Descriptive summary (low-income, pre-ACA)

	Treatment States (mean)	Control States (mean)
Age	45.82	46.09
Age 27-34	20%	20%
Age 35-42	20%	20%
Age 43-50	21%	21%
Age 51-57	20%	20%
Age 58-64	19%	19%
Male	0.43	0.41
White	0.67	0.65
Married	0.3	0.29
Household size	2.92	2.8
Number of dependent children	0.98	0.95
Have college degree	0.13	0.12
Employed	0.43	0.44
Income as FPL	75%	75%
Newly eligible	0.69	0.75
Homeownership	0.39	0.44
Having mortgage	0.24	0.24
ZHVI	$225,086	$166,000
Observation	311252	392779

Note: Data from the ACS 2009-2017. Newly eligible represents the share of the observation that would be newly eligible for Medicaid if the state expanded Medicaid. ZHVI is the Zillow median housing value index at state level, adjusted by CPI.

Table 1.2. Estimated DDD effect on homeownership

	Full	age 27-34	age 35-42	age 43-50	age 51-57	age 58-64
Basic model	-0.001	-0.003	0.001	0.002	0.000	0.033**
	(0.009)	(0.009)	(0.012)	(0.012)	(0.010)	(0.014)
Add state specific time trend	0.010	0.002	0.012	0.015	0.012	0.043***
	(0.008)	(0.010)	(0.009)	(0.009)	(0.010)	(0.015)
Flexible model	-0.003	-0.013*	-0.003	0.012	0.008	0.025
	(0.005)	(0.007)	(0.010)	(0.009)	(0.012)	(0.018)
Mean	0.41	0.23	0.34	0.42	0.49	0.56
N	1222280	241780	237454	246107	245929	251010

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Dependent variable is the homeownership status. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.3. Estimated DDD effect on having mortgage

	Full	age 27-34	age 35-42	age 43-50	age 51-57	age 58-64
Basic model	0.011	-0.002	0.011	0.017	0.022	0.043***
	(0.010)	(0.008)	(0.014)	(0.010)	(0.013)	(0.012)
Add state specific time trend	0.015*	0.002	0.018	0.022**	0.025**	0.042***
	(0.009)	(0.008)	(0.011)	(0.010)	(0.012)	(0.012)
Flexible model	0.005	-0.008	0.007	0.018*	0.011	0.023
	(0.007)	(0.006)	(0.010)	(0.010)	(0.011)	(0.014)
Mean	0.22	0.15	0.22	0.25	0.24	0.24
N	1222280	241780	237454	246107	245929	251010

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Dependent variable is mortgage status. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.4. Estimated DDD effect on outright homeownership

	Full	age 27-34	age 35-42	age 43-50	age 51-57	age 58-64
Basic model	-0.012**	-0.002	-0.010	-0.014	-0.022*	-0.012
	(0.005)	(0.004)	(0.007)	(0.009)	(0.012)	(0.011)
Add state specific time trend	-0.007	-0.000	-0.007	-0.006	-0.013	-0.001
	(0.004)	(0.005)	(0.007)	(0.007)	(0.012)	(0.010)
Flexible model	-0.007	-0.004	-0.011	-0.004	-0.002	0.000
	(0.005)	(0.004)	(0.007)	(0.008)	(0.014)	(0.011)
Mean	0.19	0.08	0.12	0.17	0.24	0.32
N	1222280	241780	237454	246107	245929	251010

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Dependent variable is outright homeownership. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.5. Estimated DDD effect on housing market outcomes (top uninsured rate)

	Annual sales units (log)	Home value (log)	New housing units (log)
Basic model	0.079	0.139**	0.032
	(0.085)	(0.070)	(0.187)
Add state specific time trend	0.079	0.140**	0.032
	(0.085)	(0.070)	(0.188)
Flexible model	0.266***	0.122	0.018
	(0.092)	(0.076)	(0.110)
N	114192	114192	23877

Notes: Data from the Zillow Research 2009-2017 are used to obtain estimates of annual sales volume and home value. Data from the Building Permits Survey 2009-2017 are used to obtain estimates of new housing units. Standard errors in parentheses are clustered at the zip code or county level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.6. Estimated DDD effect on homeownership (age 70+)

	Homeownership	Having mortgage
Basic model	0.001	-0.008
	(0.012)	(0.005)
Add state specific time trend	0.017*	-0.006
	(0.010)	(0.005)
Flexible model	0.032***	-0.013
	(0.008)	(0.008)
Mean	0.65	0.16
N	376223	376223

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Sample includes individuals older than 70. Dependent variable is homeownership. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.7. Estimated DD effect on homeownership
(individuals with family income > 138 % of FPL)

	Full	age 27-34	age 35-42	age 43-50	age 51-57	age 58-64
Basic model	0.015	0.009	0.020	0.015*	0.009*	0.024
	(0.009)	(0.006)	(0.013)	(0.008)	(0.005)	(0.019)
Add state specific time trend	-0.002	0.005	-0.006	0.000	0.002	-0.014***
	(0.002)	(0.003)	(0.004)	(0.003)	(0.002)	(0.005)
Flexible model	0.002	-0.000	0.004	0.003	0.001	0.003
	(0.003)	(0.004)	(0.004)	(0.005)	(0.003)	(0.005)
Mean	0.75	0.54	0.7	0.78	0.83	0.86
N	5118540	1159315	927481	1149969	1102667	779108

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Sample includes individuals with family income greater than 138% of FPL. Dependent variable is homeownership. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 1.8. Estimated DD effect on having mortgage
(individuals with family income > 138 % of FPL)

	Full	age 27-34	age 35-42	age 43-50	age 51-57	age 58-64
Basic model	0.004	-0.004	0.012	0.005	-0.004	0.019
	(0.005)	(0.006)	(0.011)	(0.006)	(0.007)	(0.017)
Add state specific time trend	-0.005**	0.002	-0.009*	-0.005*	0.002	-0.018***
	(0.002)	(0.004)	(0.005)	(0.003)	(0.004)	(0.006)
Flexible model	-0.001	-0.003	0.001	-0.000	-0.001	0.001
	(0.005)	(0.007)	(0.005)	(0.007)	(0.005)	(0.005)
Mean	0.58	0.48	0.62	0.64	0.61	0.53
N	5118540	1159315	927481	1149969	1102667	779108

Notes: Data from the ACS 2009-2017 are used to obtain these estimates. Sample includes individuals with family income greater than 138% of FPL. Dependent variable is homeownership. Models control for age, gender, race/ethnicity, household size, income to poverty ratio, education, marital status, and. employment status. Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Figure 1.1. Time trends of homeownership among treatment and control groups

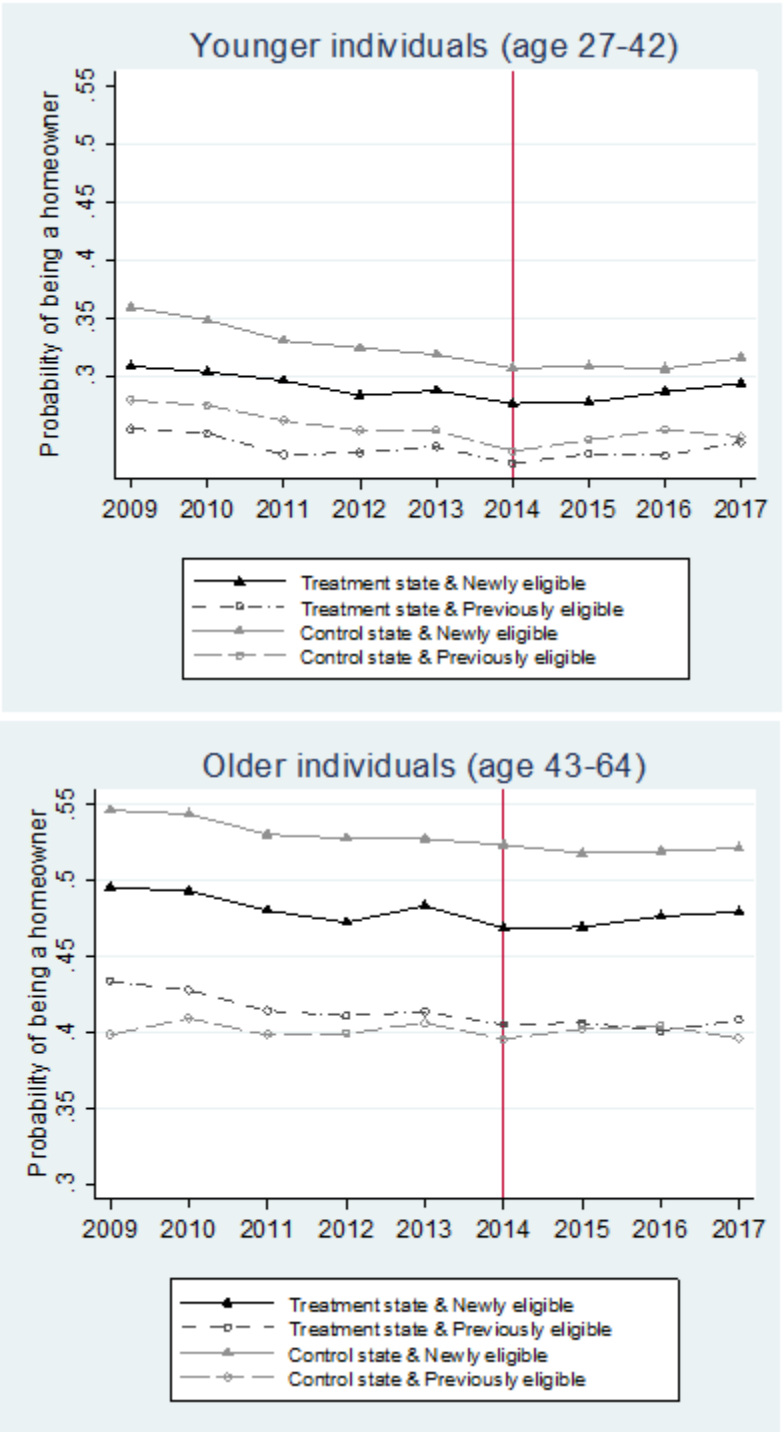


Figure 1.2. Estimated effect of the policy on homeownership (by age group)

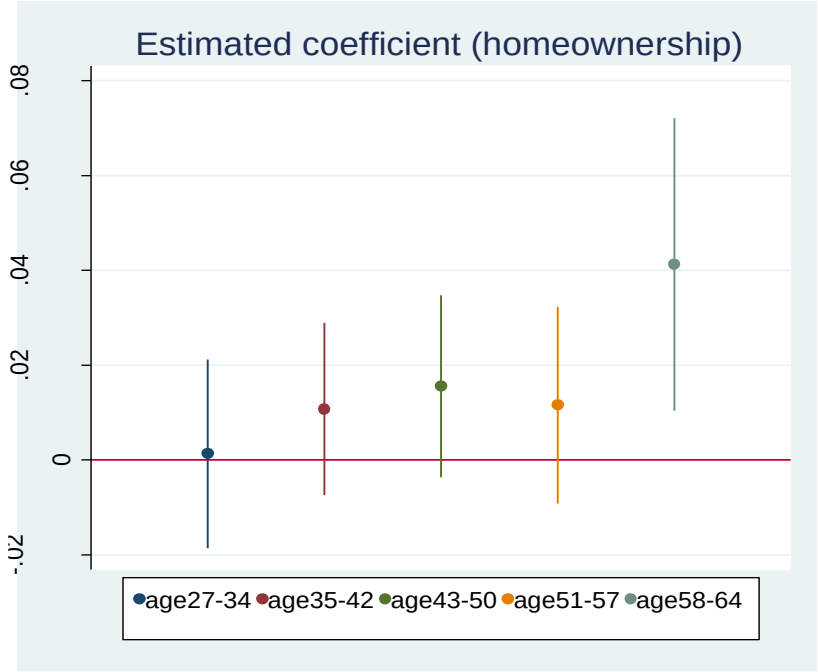


Figure 1.3. Estimated effect of the policy on mortgage status (by age group)

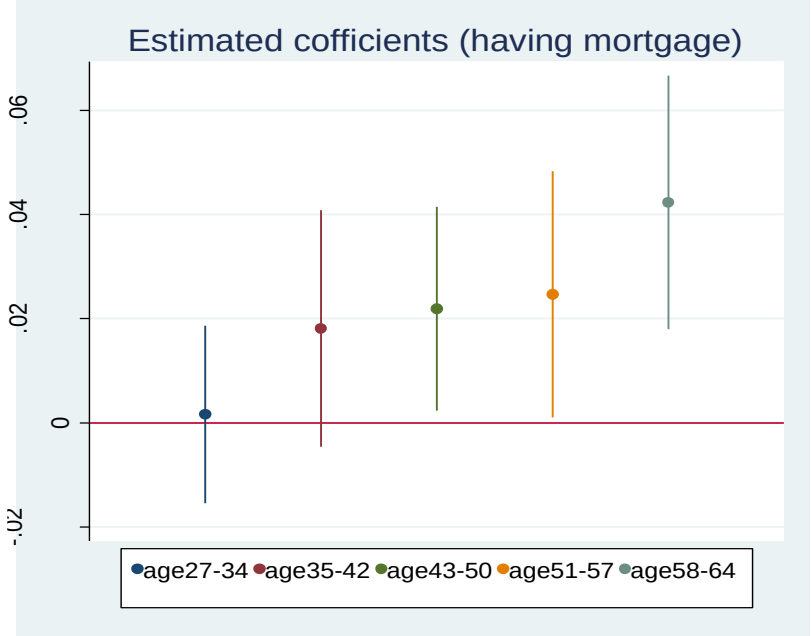


Figure 1.4. Time trends of annual sales volume among treatment and control zip codes

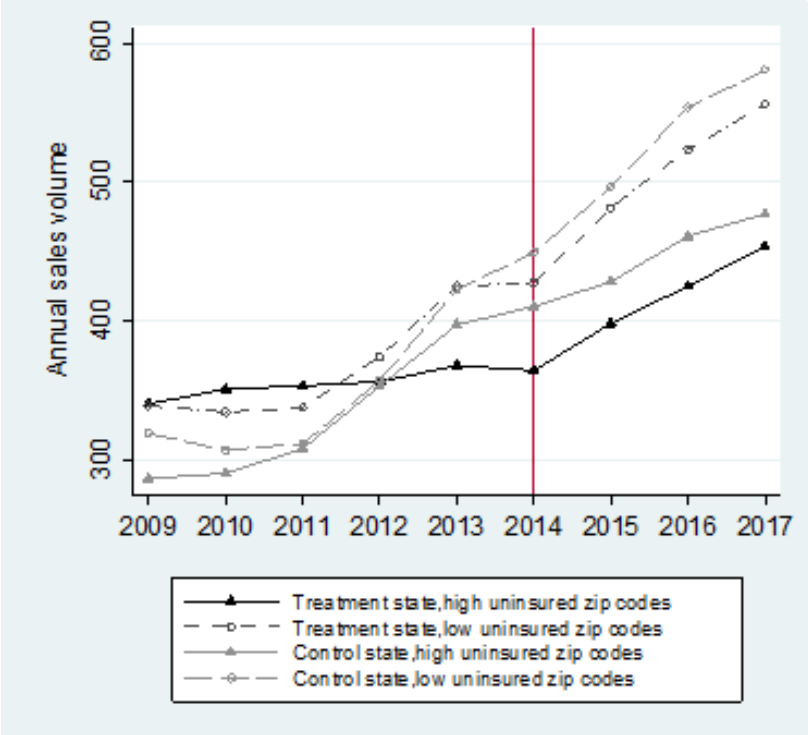


Figure 1.5. Time trends of housing price index (ZHVI) among treatment and control zip codes

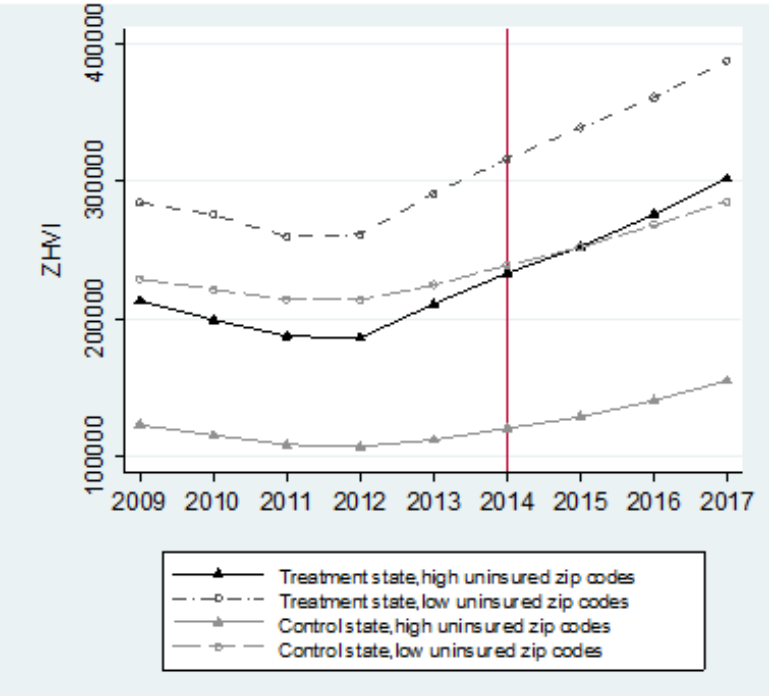


Figure 1.6. Time trends of new housing units among treatment and control counties

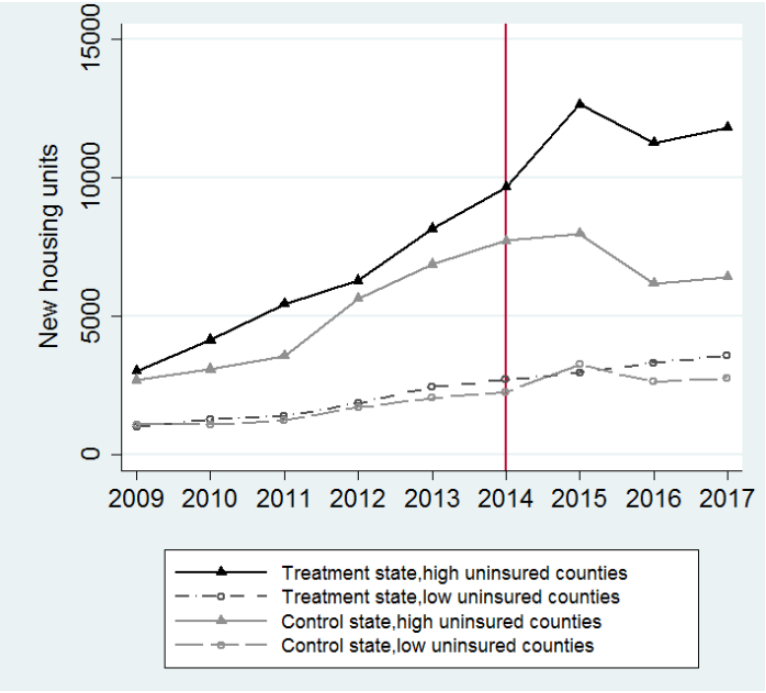


Figure 1.7. Placebo test by randomly assigning state treatment status

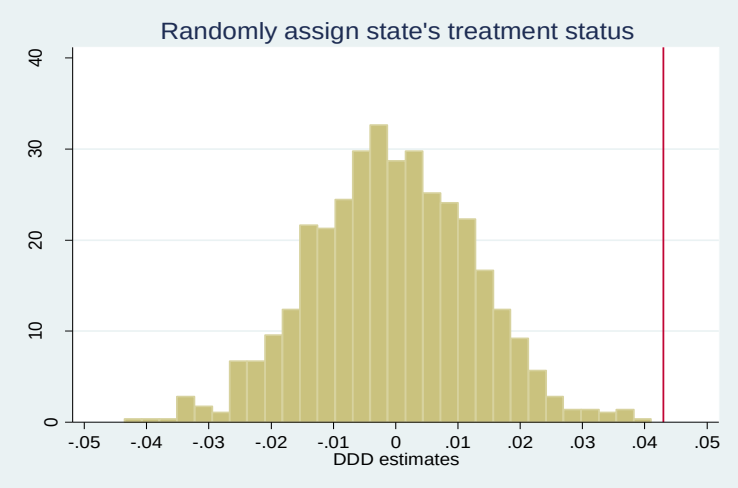
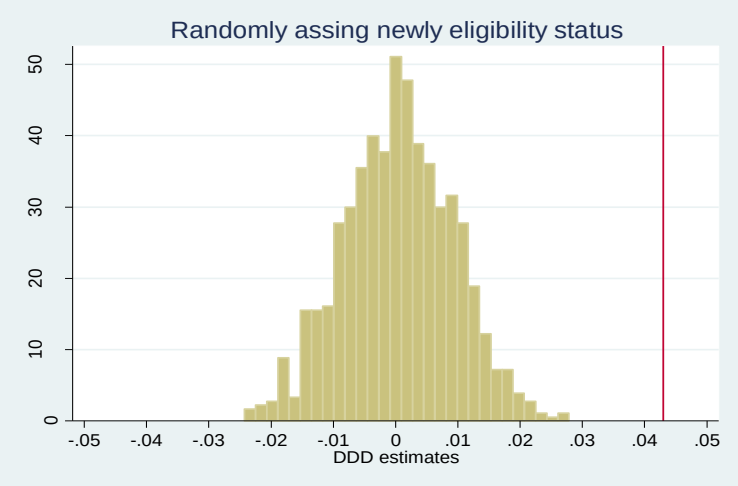


Figure 1.8. Placebo test by randomly assigning newly eligibility status



Chapter 2. The ACA Marriage Penalty and Marital Status

2.1 Introduction

Since Becker (1973, 1974), economists have understood that marriage decisions are affected by economic factors. Theoretically, individuals should choose to marry when the gains from marriage are positive. In many countries, income tax systems and social welfare programs are marriage-dependent, meaning that program features and tax treatment can affect gains from marriage and subsequent marital decisions. In the US context, an extensive literature has confirmed that marriage decisions are impacted by the structure of the U.S. income tax system (e.g., Alm and Whittington 1996; Dickert-Conlin 1998; Eissa and Hoynes 2000; Fisher 2003), and by government transfer programs such as Medicaid and Medicare, the Supplemental Nutrition Assistance Program (SNAP), the Earned Income Tax Credit (EITC), and Temporary Assistance for Needy Families (TANF) (e.g., Yelowitz 1998; Abramowitz 2015; Wilcox et al 2016; Chen 2018; Michelmore 2018). The Affordable Care Act (ACA) is the largest new tax and transfer program in the U.S. in recent years (Kucko et al., 2017), and like existing programs, confers benefits based on family and earnings structure. As such, the ACA may produce marriage penalties or subsidies, which may deter or encourage people to wed. Indeed, in the popular press the debate of whether the ACA has deterred marriage was heated. For example, an

article from The Atlantic quoted: *"The first time I heard Nona Willis Aronowitz talk about getting divorced to save money on health insurance I thought she couldn't really be serious… But with the arrival of the Affordable Care Act's insurance exchanges, the question for Nona and her husband Aaron Cassara moved from the realm of casual conversation to a real financial conundrum."*[23] However, few academic study has considered whether the main provisions of the ACA – the Medicaid expansion and the premium tax credit program – have encouraged or deterred marriage. Our goal in this study is to empirically examine for whom the ACA has generated marriage penalties or subsidies, and whether these subsidies and penalties have had an effect on marital status.

Understanding how large-scale public policies like the ACA affect household composition is important. The marriage rate has been declining for decades in the U.S., especially among people with lower social economic status (SES). More people are choosing non-marital cohabitation (Parker and Stepler, 2017). Formal marriage has been linked with family well-being, including health, wealth, and child outcomes (e.g., Waite and Gallagher, 2000). Access to certain benefits such as pensions may be available only to spouses rather than cohabitating partners. Given the fact that the ACA has affected millions of people through health insurance expansion, its spillover effect on marital status might also be substantial. The ACA's focus on lower income people also suggests that it could disproportionately advantage certain couples over others in terms of access to the benefits of marriage.

[23] See article from https://www.theatlantic.com/politics/archive/2013/11/the-hidden-marriage-penalty-in-obamacare/280890/

The ACA might affect marital decisions because program eligibility and benefit amounts are determined by benchmarking adjusted household income against the Federal Poverty Level (FPL). Under the ACA, low-income Americans with family income below 138% of the FPL[24] are eligible for Medicaid[25], which is free public health insurance with low cost-sharing. Additionally, low-to-moderate income Americans with family income between 100-400% of the FPL, who have no access to affordable employer sponsored health insurance (ESI), can receive the ACA premium tax credits to subsidize non-group private health insurance.[26] Importantly, the FPL is a household-size-specific income level, which increases nonlinearly with household size. Thus, marriage can greatly affect access to ACA benefits, since marriage changes family size – by adding a spouse and any dependent children – and household income – by adding the new spouse's earnings, if any. Hence, the ratio of household income to the FPL could be significantly altered by marriage, which may produce positive or negative economic benefits under the ACA. We call these positive or negative economic benefits *marriage penalties* henceforth, with the understanding that the marriage penalty may be negative (and therefore a subsidy).

The ACA will generate marriage penalties if marriage reduces the total ACA benefits available to a couple. For example, two single individuals, each with earnings just below the one-

[24] FPL is determined by the Department of Health and Human Services (HHS) in each year. It represents the minimum amount of income for living and is commonly used to determine the eligibility for transfer programs. FPL varied by household size.

[25] Not all states adopted Medicaid expansion. The income eligibility cutoffs were different across states before the ACA Medicaid expansion. The cutoffs remained unchanged in Medicaid non-expansion states.

[26] The amount of premium tax credits is the difference between the original premium of the second-lowest-cost silver plan and the contribution of individuals. The contribution is based on a sliding scale percentage of family income. Basically, the contribution increases with the ratio of income to the FPL, and the percentage is capped at 9.5% for those reach 400% of the FPL. For example, Families with income up to 133% of the FPL will contribute 2% of income, 3-4% for individuals with income between 133-150% of the FPL, etc. The percentage is capped at 9.5% for individual with income between 350-400% of the FPL.

person 138% FPL Medicaid-eligibility cutoff will both lose Medicaid eligibility if they wed (since their combined earnings will fall above the two-person 138% FPL cutoff). In this case, they will face a marriage penalty under the ACA because marriage will cause them to lose free public health insurance. Similarly, two single individuals who are independently eligible for premium tax credits may receive a lower subsidy if they wed, since marriage will increase their household earnings relative to the two-person FPL. On the other hand, the ACA offers a marriage subsidy to couples whose health insurance benefits will increase if they wed. For instance, an individual whose earnings place her just above the 138% FPL cutoff for one-person families stands to gain Medicaid eligibility if she marries a low-earner (since their combined earnings will leave them lower than 138% of the two-person FPL threshold). If her potential spouse has dependent children, marriage will grant her Medicaid eligibility even if her earnings are quite high relative to the one-person FPL cutoff. In sum, whether the ACA generates marriage penalties or subsidies is determined by couple-specific characteristics: individual levels of – and the combination of – income, the number of dependent children, as well as the couple's state of residence, and their ages (factors which we will discuss the detail in the next section).

We use individual-level data from the 2011 to 2017 American Community Survey (ACS), which includes a wealth of information about individual and household characteristics and geographic location. We define the ACA marriage penalty or subsidy as the difference between the total premium payment that two individuals would face if they marry, and the total payment if they are unmarried.[27] Medicaid premiums are set to zero, and age-adjusted private-

[27] We ignore any changes to copays or deductibles that may result from marriage, since these depend on healthcare usage and vary greatly across couples.

market insurance premiums are obtained from a collection of state-level Marketplace premiums published by Kaiser Family Foundation (KFF).[28] We are able to directly calculate penalties and subsides for cohabitating and married couples, since the ACS includes information on all household members. Non-cohabitating partners, however, appear as single in the ACS. For these couples, we follow the literature and simulate a marriage market, which allows us to form hypothetical couples from the single ACS respondents (Michelmore, 2018). The resulting data is a set of married and unmarried couples – either observed or simulated – for whom we can calculate marriage subsidies and penalties under ACA.

After identifying the subsidy or penalty that a couple faces, we then explore whether they respond to the ACA marriage incentives in their decision to marry. Theory predicts that as a penalty (subsidy) grows, a couple should become less (more) likely to marry. We test this hypothesis using a difference-in-differences (DD) framework that compares the marital status of couples facing different ACA marriage penalties or subsidies, before and after the ACA became effective. For example, take a couple facing a $1,000 marriage penalty after the enactment of the ACA in 2014. We compare this couple's marital status to that of a similar couple who would have faced the same penalty had the ACA policy been in place before 2014, and to that of a couple with a smaller penalty post-2014. Our approach seeks to identify differences in marital status between these three couples, each of which faces different policy-induced financial gains of marriage.

[28] Second-lowest cost silver plan is the benchmark plan that determines the amount of premium tax credits. Premium of the plan varies by age, location, and tobacco use. Pre-existing health conditions no longer affect premium under the ACA.

We also consider the possibility that marriage penalties and subsidies are correlated with changes in characteristics that also affect marriage decisions. For example, if both partners work and earn higher incomes, they are likely to face a high ACA marriage penalty. In this case, since socioeconomic status is a correlate of marriage, we may spuriously conclude that higher marriage penalty is associated with higher probability of being married. More worrisome is the possibility that couples adjust income in response to the policy change, something that is likely given the extent to which health insurance is tied to work. In this case, our empirical approach may misidentify the relationship between ACA marriage incentives and marital status. To address these potentials sources of bias, we additionally use a two-stage simulated instrument variables (IV) approach. We first simulate the couples' exposure to ACA marriage penalties or subsidies based on their state of residence, age, and number of dependent children. This measure of exposure is independent of income, and variations are mainly driven by exogenous policy features. We predict the couples' actual marriage penalties or subsidies using the exposure measure. In the second stage, we use the predicted marriage penalties or subsidies on marital status to see if the probability of being married is different across individuals facing different marriage penalties or subsidies before and after the ACA went into effect.

We find that the ACA marriage incentives affect the probability of marriage. The probability of being married drops by 0.7 percentage points with a $1,000 increase in the ACA marriage penalty. However, our IV estimates are much larger in magnitude: the probability of being married drops by 6 percentage points with a $1,000 increase in the ACA marriage penalty – an effect that translates into a 9 percent reduction in marital rates. Our results are robust to different ways of calculating the ACA marriage penalty. In addition, our results also show that

the ACA had heterogenous effects among groups with different SES. Marriage penalties had larger effects on individuals who are less educated. A $1,000 increase in the marriage penalty reduced the probability of being married by 7.25 percentage points (12.4 percent) among those without a college degree, while it only reduced the probability of being married by 2.87 percentage points (3.9 percent) among those with college degrees. Racial/ethnic groups other than White are also more responsive to the ACA marriage penalties or subsidies. A $1,000 increase in the marriage penalty reduced the probability of being married by 8.6 percentage points (16.8 percent) among racial/ethnic groups other than White, while it only reduced the probability of being married by 5.3 percentage points (7.9 percent) among White people.

This study contributes to the literature about the social impacts of the ACA health insurance expansion. Most empirical studies on the ACA focus on the effect of policy on health insurance coverage and healthcare utilization (e.g., Frean et al. 2017; Simon et al. 2017; Courtemanche et al. 2017). A growing body of research explores its impact on individual financial health and labor supply, as well (e.g., Busch et al. 2014; Hu et al. 2016; Kaestner et al. 2017; Leung and Mas, 2016). In terms of household composition, a small group of studies examine the effects of health insurance expansion on marital status. However, these studies mainly focus on the fact that health insurance coverage through the ACA dependent mandate, Medicare and Medicaid programs increase access to alternative affordable health insurance schemes outside of marriage and employment, and thus reduce the incentive of being married (e.g., Abramowitz, 2015; Chen, 2018; Slusky and Ginther, 2018; Hampton and Lenhart, 2019). Our study looks at a similar question from a different angle. Similar to the marriage penalties or subsidies embedded in the U.S. personal income tax system, the ACA eligibility and generosity

structure produces heterogeneity in marriage incentives. Thus, we add to the existing literature by quantifying the amount of penalties or subsidies that people face under the ACA. Using these values, we are able to provide estimates of the elasticity of marriage with respect to income.

2.2 Background

2.2.1 Taxation and Marriage

Reforms in the tax and transfer policy generate behavioral responses, such as changes in labor supply and marriage decisions. The link between taxation and marriage mainly come from the policy influences on the costs and benefits of marriage. Becker's theoretical model of marriage (1973, 1974) suggests that individuals will choose marriage when the total output of marriage exceeds the sum of the outputs of each individual if remained unmarried. Output can be broadly defined, such as home production (the quality of meals, the quality and quantity of children, companionship and love, prestige, and health status) and welfare or other economic gains and losses.

Based on Becker's model, an extensive body of studies have examined the effects of the U.S. income tax system on marriage (e.g., Alm and Whittington 1996; Dickert-Conlin 1998; Eissa and Hoynes 2000; Fisher 2003). The U.S. income tax system is not neutral towards marriage because tax brackets differ between married filing jointly and singles.[29] In other words, total tax liability for the married couple could be different from the total tax liability of the two if

[29] This is the case before 2002, where the tax bracket threshold for married filing jointly was less than twice the threshold for singles in all tax brackets. Since 2002, the U.S. income tax system becomes more neutral towards marriage, with the bottom two tax brackets started to be equal. Since 2018, the bottom five tax brackets started to be equal. https://files.taxfoundation.org/legacy/docs/fed_individual_rate_history_nominal.pdf

they were unmarried. For example, marriage might generate penalties when marriage makes combined income fall into a higher tax bracket. This happens in the cases where couples earn similar income. However, marriage might generate subsidies when one spouse earns much less than the other, and thus moves combined income into a lower tax bracket (see Pomerleau 2015 for more detailed examples).

Studies of income taxes and marriage define marriage penalties or subsidies as the difference between the tax liability of a couple if married and the sum of tax liability of each partner if they remained unmarried. Alm and Whittington (1998) find that cohabiting couple are less likely to transit to formal marriage when facing larger marriage penalties. Eissa and Hoynes (2000) shows that the probability of being married dropped by 0.4 percentage points with a $1,000 increase in marriage penalties. Using similar methods and data, Fisher (2013) shows that the probability of being married reduced by 1.7 percentage points with a $1,000 increase in the marriage penalty, after adjusting for the potential endogeneity between marriage penalty and marital status.

Studies have also considered the effect of the Earned Income Tax Credit (EITC) on marital status. The schedule of the EITC, which is based on household income but does not increase linearly for married couples, creates a marriage disincentive for certain couples. Ellwood (2000) finds positive but modest effect of the EITC expansion on marriage. Michelmore (2018) shows that the likelihood of getting married among single mothers would decrease by 2.5 percentage points if they expect to lose EITC eligibility because of marriage.

2.2.2 Health Insurance Coverage and Marriage

A growing literature has investigated the link between health insurance and marriage choice. In the U.S., health insurance coverage is closely tied to employment. A vast majority of nonelderly adults are covered by employer sponsored health insurance (ESI). For individuals who have no access to ESI and are not eligible for public health insurance such as Medicaid, it is difficult to obtain health insurance because private individual health insurance is very expensive. Thus, marrying a partner with ESI that offers spousal coverage increases the benefits of the marriage. Anecdotal evidence from media reporting and surveys suggests that people do choose to marry to access health insurance (e.g., Kaiser Family Foundation, 2008; Goodman, 2008).

Studies also have shown that when government transfer programs facilitate obtaining affordable health insurance coverage outside of marriage and employment, marriage rates decrease. Abramowitz (2015) shows that the ACA dependent mandate, which allows younger adults (age 19-26) to stay in their parents' private health insurance plan, reduces the probability of marrying by about 0.5 percentage points (9 percent) and increases the probability of divorce by 0.3 percentage points (10 percent). Chen (2018) explores the effect of being eligible for Medicare on marriage choice among older individuals. Being eligible for Medicare increased the likelihood of getting divorced by about 7 percent for those who had spousal health insurance compared to those who did not. Hampton and Lenhart (2019) investigate the influence of the ACA Medicaid expansion on marriage. They find that increased access to Medicaid reduced the probability of being married by 2.13 percent and increased the probability of being divorced by 3.82 percent. Wilcox et al (2016) investigates the impact of Medicaid, TANF, and SNAP on marital status among couples with young children. They find that the marriage penalty in

Medicaid eligibility or food stamps reduce the probability of being married by about 2 to 4 percentage points.

2.2.3 Linking the ACA and marriage

The Affordable Care Act (ACA), which was enacted in 2014, is the largest tax and transfer program in the U.S. in recent years (Kucko et al., 2017). The ACA aims to provide free and subsidized health insurance to low and low-to-moderate income Americans. There are two main mechanisms by which the ACA provides subsidized health insurance: Medicaid – which was expanded under the ACA – and the premium tax credit programs – which allows for purchase of subsidized insurance through the Health Insurance Exchanges (or Marketplaces). As of 2016, about 20 million uninsured nonelderly Americans gained health insurance under the ACA (Kaiser Family Foundation, 2016). According to Frean et al. (2017), Medicaid expansion accounted for 60 percent of the coverage gains, and the Health Insurance Exchanges accounted for 40 percent of the coverage gains.

The Medicaid expansion extended free public health insurance to low income Americans. Before the ACA, income eligibility cutoffs for Medicaid differed across states, and were generally very low. Childless adults were also usually ineligible for Medicaid. In 2014 when the ACA went into effect, the federal government offered funding to states to allow them to increase the income eligibility cutoffs to 138% of the Federal Poverty Level (FPL), and to eliminate the parental requirement. Nineteen states, however, did not accept the funding and chose to keep their eligibility cutoffs at the pre-ACA level, and in most cases, exempt childless adults from coverage. Between 2014 and 2017, seven more states chose to expand Medicaid. Table A.1

summarizes the ACA income eligibility cutoffs and adoption status of Medicaid expansion across states.

The premium tax credits are available to individuals with family income between 100-400% of the FPL, and who have no access to affordable employer sponsored health insurance (ESI).[30,31] The value of the credits are determined by a schedule that defines the total contribution that a family should make towards their insurance premiums. Table B.1 presents the contribution schedule for the ACA premium tax credits. In essence, the total allowable contribution that a family must pay increases with the ratio of income to the FPL, from 2% of income for families earning 100% FPL to 9.5% of income for those earning 400% FPL.[32] The value of the credit depends on the unsubsidized premium a family faces – which the government defines as the premium the family would have to pay if they bought the second lowest-cost silver plan on a Health Insurance Exchange. This premium, in turn, can depend on factors like age and state of residence. The amount of premium tax credits is substantial, advanceable, and refundable. To be more specific, premium tax credits can be received in advance to pay for health insurance premium. The amount is determined by the projected annual income. Premium tax credits then will be reconciled based on actual income when people file tax return in the following year. People will either get refund or pay the difference between the actual subsidies they are eligible for and the subsidies they already received in the previous year.

[30] In Medicaid expansion states, the eligibility threshold for premium tax credits starts at 138% of the FPL.
[31] There is a "coverage gap" in Medicaid non-expansion states, where a group of individuals earn too much to be eligible for Medicaid and too little to be eligible for premium tax credits.
[32] It could be possible that the premium payment based on the ACA contribution schedule is larger than the original health insurance premium. In this case, those individuals receive no premium tax credits. For example, younger adults face much lower insurance premium than older individuals. If they earn a relatively high income, say 399% of the FPL, they would have to pay 9.5% of their income, which could be greater than the original health insurance premium for them.

In the context of the ACA, we define a marriage penalty/subsidy as the difference between the family's total health insurance premium payment if the couple marries, and the sum of each individual's premium payment if they are unmarried. In both cases, these include any premiums for dependent children that either partner must pay. If they are eligible for Medicaid, we set the premium at zero. The specifics of the ACA eligibility and generosity rules described above generate marriage penalties/subsidies that differ across couples along several dimensions, including income of each partner, state of residence, age, and number of children. The remaining part of this section illustrates the sources of variation in more detail.

Income

Most crucially, program eligibility and generosity are determined by the ratio of household income to the household size-specific FPL. The Federal Poverty Level income threshold increases nonlinearly with family size.[33] For instance, the 2014 FPL for a two-person family was $15,730, far less than two times the one-person income threshold of $11,670. When two individuals marry, both their household size and family income change; the ratio of their adjusted income to the FPL changes; and thus the eligibility of Medicaid or premium tax credits, and premium tax credits are affected. This, in turn, leads to marriage penalties that depend on each partner's earnings, as well as the couple's joint earnings.

[33] For example, the FPL for a single person family is $11,670 in 2014. The FPL increased by $4,060 for each additional person. Thus, the FPL for a family of two is $15,730 and the FPL for a family of three is $19,790, etc. A single person earns $20,000 a year would have the ratio of income to the FPL equals 171% ($20,000/$11,670), a family of two with of $20,000 /year would have the ratio of income to the FPL equals 127% ($20,000/$15,730) and a family of three with 20,000 income would have the ratio of income to the FPL equals 101% ($20,000/$19,790).

For example, take an unmarried couple of two childless people, who each earn $20,000 annually, and each have a household size of one. If they remain unmarried, they each face a ratio of income to the FPL equal to 171% ($20,000/$11,670). Now suppose that the couple marries: their combined income would be $40,000, and their family size has increased to two. The ratio of their adjusted income to the FPL as a married couple is equal to 254% ($40,000/$15,730). This puts them in a higher contribution bracket, leaving them eligible for a less generous tax credit as a result of the marriage. This couple would therefore experience a marriage penalty.

Panel A in Table B.2 provides some hypothetical scenarios of marriage penalties for a couple composed of two 30-year-old childless single adults living in a Medicaid expansion state. The first four cases show various marriage penalties if individual A earns $10,000, and is therefore Medicaid eligible before marriage. The chart shows how the couple's marriage penalty changes for different earnings levels of individual B. If person B also earns $10,000, there is no penalty since both partners are Medicaid-eligible regardless of marital status. However, as person B's earnings increase, the marriage penalty changes in unexpected ways. If Person B earns $30,000 the penalty grows to $850; if they earn $50,000, the penalty shrinks into a subsidy of $300, since Person B gains access to the premium tax credit through marriage; and if person B earns $60,000, the marriage penalty grows to $6,000, since marriage leaves the couple ineligible for any subsidized ACA program. Further, comparing cases 2 vs. 5, and cases 3 vs. 6 shows that, even when combined income is the same, the penalty varies as the distribution of income across partners changes.

Figure B.2 plots simulated marriage penalties for a hypothetical couple where both partners are 30 years old and live in a Medicaid expansion state. The figure, which is a heat map,

shows how the predicted penalty changes as each partner's income varies from zero dollars to 50,000 dollars per year. The figure demonstrates that if both partners earn relatively little, there is no marriage penalty. As one or the other's income increases, penalties start to increase: once one partner earns at least 20,000 dollars per year, the couple are at risk of experiencing a marriage penalty, a risk that increases as the couple's total income increases. The figure also shows how couples with equal levels of total income can face very different marriage penalties depending on how that income is distributed across partners.

State of residence

ACA benefit eligibility – and therefore the ACA's marriage penalty – is also affected by a couple's state of residence in two ways. First, Medicaid eligibility rules are state-specific because of the fact that some states did not expand Medicaid. In Medicaid non-expansion states, individuals will face a coverage gap if their earnings fall above the state's Medicaid eligibility threshold, but below 100% FPL where the premium tax credits kick in. This creates a very different set of marriage premiums than those facing couples living in Medicaid expansion states.

Panel B in Table B.2 illustrates this difference. The panel shows hypothetical scenarios of marriage penalties for the same 30-year-old childless couple, who we now assume live in a non-expansion state. Case 7 of the table illustrates the marriage penalty if both partners earn $10,000 dollars per year (86% of the one-person FPL). Whereas in an expansion state this couple would face no marriage subsidy (case 1), in a non-expansion state, their subsidy grows to $11,600. This is because when unmarried, each partner's earnings leave them in the coverage gap; when married, the couple's joint earnings of 20,000 dollars leave them at 127% of FPL

($20,000/$15,730), and give them access to a sizeable premium tax credit. Thus, similar couples living in Medicaid expansion and non-expansion states face different marriage penalties. This difference only applies to low-income individuals, however, who would be eligible for Medicaid if all states had expanded Medicaid. For individuals with higher incomes who are ineligible for Medicaid regardless of state, there is no difference in marriage penalty between Medicaid expansion and non-expansion states (comparing cases 5 and 11), assuming original insurance premiums are the same across states (simplified case). Figure B.2 shows another heat map that plots the simulated marriage penalties for a 30-year-old childless couple that lives in a Medicaid non-expansion state. The figure is similar to B.1, except it now includes large marriage subsidies for low earning couples who fall into the coverage gap when unmarried.

Under the ACA, health insurance premiums can vary by area of residence. Thus, the second way that state of residence affects the marriage penalty is through state-level differences in Health Insurance Exchange premiums. Since the value of a couple's premium tax credit is equal to the difference between their unsubsidized insurance premium and their contribution limit (9.5 percent of total earnings, for example), the value of the tax credit will vary by state.

Number of children

The number of dependent children in a couple also affects the marriage penalty because it affects the household size used to determine the FPL. For families not eligible for any ACA benefits, total premium payment is the sum of original premium of each family member. Table B.3 panel A shows how the ACA marriage penalty changes with number of children for a hypothetical 30-year-old couple where both partners earn 30,000 dollars per year. Adding

children to partner B affects their unmarried FPL ratio – and therefore their premium tax credit; it also affects the married couple's FPL ratio. As a result, the couple's marriage penalty changes non-monotonically with number of children. The change is even larger if we increase person A's earnings to 50,000 dollars per year (Panel B). In this case, the penalty grows from $3,570 to $6,110 with the addition of the first child; it then shrinks to only $400 with the addition of a second child. Figure B.3 shows the heat map that illustrates how the marriage penalty changes with the number of dependent children. For simplicity, we only show the simulated results assuming the couple live in a Medicaid non-expansion state. Premium payments for children are set to zero if the family falls in the coverage gap, because those children are eligible for children's health insurance program (CHIP).

Age

Under the ACA, insurers are allowed to charge different premiums based on applicants' ages. Health insurance premiums are generally higher for older individuals, up to a maximum cap. The base insurance premium is calculated using a 21-year-old policyholder. Then an adjustment factor is used to calculate the insurance premium for a specific age. For example, the adjustment factor is 0.83 for a 15-year-old child, 1.14 for a 30-year-old adult, 1.79 for a 50-year-old adult. The adjustment factor is capped at 3 when an individual reaches 64-year-old. Thus, the marriage penalties or subsidies are different for otherwise similar couples who have different age combinations.

To sum up, the total amount of premium payments may be different when marital status changes, and the change of premium payments varies by income, state of residence, number of

children, and age. Hence, the ACA generates marriage penalties or subsidies that depend on a complex interaction of these factors. In the following section, we describe the data and empirical strategy used to examine the impact of these ACA marriage penalties on marital status.

2.3 Data

We use the American Community Survey (ACS) 2011 to 2017.[34] The ACS is the largest U.S. cross-sectional household survey that interviews about 3 million individuals every year. The dataset contains rich information about individual characteristics including incomes, marital status, fertility, living arrangement, and state of residence.

The ACS surveys are conducted at the household level, where all individuals living in the same house are observed. With the available information about relationships between all individuals in the same house, an individual can be identified as married with spouse present, unmarried cohabitant, or single. Our identification strategy requires observation of income, age and number of children for both partners. For married couples with spouse present, and for unmarried cohabitants, both partners' characteristics are observed. For individuals who live alone, however, any potential partner's characteristics are not observed. We therefore simulate a marriage market for single individuals (details are discussed below) to predict their potential partner's characteristics. We restrict the sample to individuals who are most likely to be affected by the ACA health insurance expansion and thus face embedded marriage penalties or subsidies.

[34] Ruggles S, Flood S, Goeken R, et al. IPUMS USA: Version 7.0 [dataset]. Minneapolis, MN: Integrated Public Use Microdata Series; 2017.

To be more specific, we limit the sample to individuals between the ages of 26 and 64[35] who meet the following conditions: either the individual, or an observable partner, has income below 400% of the FPL; for married and cohabitating people, their partners are between the ages of 26 and 64; and combined income of the household is less than $250,000. The final analytical sample includes 6,778,359 individuals. Of these, 4,320,363 are married, 457,591 are unmarried cohabitants, and 2,000,405 are singles.

Table 2.1 summarizes the demographic characteristics of individuals in the analytical sample. Unmarried cohabitants are relatively younger than married or single people. Compared to married people, unmarried cohabitants and singles are less likely to be White, less educated, earned less, and have fewer dependent children. Married men have higher income than their cohabiting and single counterparts, while earnings do not vary greatly across marital status for women.

The main dependent variable of interest is whether the couple is married or not at the time of the survey. The main independent variable is the ACA marriage penalty or subsidy that the couple faces. We calculate each couple's marriage penalty by subtracting their total premium payments if unmarried from their total premium payments if married. The ACS does not provide information about health insurance premiums that individuals face. As such, we estimate each couple's health insurance premium. For individuals that are income eligible for Medicaid, we set their premium to zero. For those who are ineligible for Medicaid, we estimate their premiums. For post-ACA years 2014 through 2017, we use data from the Kaiser Family Foundation about

[35] Individuals over 65 are eligible for Medicare and young adults up to age 26 can be included in their parents' private health insurance under the ACA.

the Marketplace Average Benchmark Premiums for each state. [36] The benchmark premiums present the second-lowest-cost silver premium for an average 40-year-old individual in each year and state. Under the ACA, insurance companies can only set insurance premium based on five things: age, location, tobacco use, individual vs. family enrollment, and plan category (Bronze, Silver, Gold, Platinum, and Catastrophic). We use the Federal default standard age curve to adjust the health insurance premium for differently aged individuals. [37] Following Frean et al. (2017), we assume all individuals eligible for premium tax credits would purchase a Silver plan because it is the most commonly purchased plan in the Marketplace. We do not adjust the premium based on tobacco use because of data limitations in the ACS. For pre-ACA years 2011 to 2013, we use reports from the Office of the Assistant Secretary for Planning and Evaluation (ASPE) and Kaiser Family Foundation (KFF) to impute the insurance premiums.[38]

2.4 Empirical strategy

2.4.1 Simulated Marriage Market

Because we cannot observe a partner's of individuals in our sample who live alone, we simulate a marriage market and match the singles in our sample to likely partners. We are then able to compute marriage penalties for these simulated couples, as well as the married and cohabitating

[36] https://www.kff.org/health-reform/state-indicator/marketplace-average-benchmark-premiums/?currentTimeframe=0&sortModel=%7B%22colId%22:%22Location%22,%22sort%22:%22asc%22%7D

[37] https://www.cms.gov/CCIIO/Resources/Regulations-and-Guidance/Downloads/Final-Guidance-Regarding-Age-Curves-and-State-Reporting-12-16-16.pdf

[38] The report from ASPE includes a comparison of individual market premiums between 2013 and 2017 in 39 state. For missing states, we assume individual insurance premium doubled from 2013 to 2017, because the individual insurance premium in those available states doubled on average at this time period. After imputing insurance premium in 2013, we assume annual growth of premium from 2011 to 2013 is 5% based on the historical national trend of individual insurance premium presented on https://www.kff.org/private-insurance/issue-brief/individual-insurance-market-performance-in-2018/.

couples in our sample. Following Michelmore (2018), we construct a marriage market for single individuals based on their race/ethnicity, age, and education. We divide all single individuals into four race/ethnicity groups: White, Black, Hispanic, Asian/other; five age groups: 26-33, 34-41, 42-49, 50-57, 58-64; and two education groups: less than college degree and have college degree. Single individuals are then randomly matched to a single opposite-sex individual within the same race-age-education cell. Since the ACS has a very large sample size, the size of each race-age-education cell is large, providing significant variation in potential partners; income, age, and number of children. Table B.4 shows the average income, age, and number of children of predicted partners.

2.4.2 Calculation of marriage penalties

Marriage penalties are defined as the difference between the total out-of-pocket health insurance premium payment if married and the total payment if unmarried. For each couple in our sample[39], we calculate this regardless of whether we observe them as married or unmarried by determine the couple's total premium payment both if they were married and unmarried. For couples we observe in 2014 and later, these are the actual marriage penalties they face; for couples in 2011 through 2013, these are the penalties they *would* face if the ACA had been in place in the year we observe them.

For example, take a couple whose joint income leaves them below the state's Medicaid income eligibility cutoff: their married premium payment is zero because Medicaid does not require a premium. If their joint income is between 100-400% of the FPL and thus they are

[39] Same-sex couples are excluded in this study.

eligible for the premium tax credit, their married premium payment equals a percentage of their family income (based on the contribution schedule shown in Table B.1). If they are eligible for neither Medicaid nor premium tax credits, their married premium payment equals the original insurance premium for a family, which is the sum of the insurance premium of each adult and child.[40] We then take the same couple and calculate each partner's individual premium, and sum them. This is the couple's unmarried premium payment. Their marriage penalty is the difference between these two values.

One complication that arises in calculating the unmarried premium is determining which partner has custody of any dependent children (since out-of-pocket premiums depend on the number of dependent children in each partner's household). For unmarried couples in our sample, this is simple since we are able to match any dependent children to the likely claiming partner. However, for couples whom we observe as married, it is not clear which partner would claim any dependent children when we calculate their unmarried premium. Following previous literature (Eissa and Hoynes, 2000; Fisher, 2013), we assume that the wife would have custody of all children if the couple were unmarried. However, as discussed by Alm and Whittington (1996), different assignments of children can produce very different estimate of marriage penalties. For a sensitivity check, we also calculate two additional marriage penalties: by assigning children to the husband, and by assigning them to the higher earner.

Figure B.4 shows the distributions of estimated ACA marriage penalties or subsidies using different assumptions about the assignment of children if married couples were unmarried. The figures show that marriage penalties are smaller when we assign children to men, or to the

[40] Individual premium payment for a child is set to zero if the child is eligible for CHIP.

higher earner. This is because women tend to be the lower earners in opposite-sex partnerships, and thus are more likely to be eligible for Medicaid or get larger premium tax credits regardless of having dependent children. Men usually earn more and are therefore eligible for less generous ACA premiums. Assigning children to men, therefore, has a larger impact on their ACA eligibility or premium payments. Thus, the couple would face smaller marriage penalties when assigning children to men or the higher earner if they divorce. We conduct our analyses using these different assumptions as a robustness check.

Table 2.1 shows that the average marriage penalty in our analytical sample is about $1,680 for married individuals, $483 for unmarried cohabitants, and $784 for single individuals whom we have matched in our simulated marriage market. The proportion of married individuals facing marriage penalties is 69%, which is higher than unmarried cohabitants (58%) and single individuals (46%). Figure 2.1 shows the distribution of the marriage penalties or subsidies. The majority of the sample (98%) faces between -$8,000 (subsidies) and $8,000 (penalties).

Figures 2.3 through 2.6 present heat maps similar to the ones discussed above. Whereas the figures in the B.1 through B.3 use hypothetical couples of different earnings levels, capturing only the policy variation in ACA marriage penalties, the figures below represent the actual distribution of penalties faced by couples in our sample. The variation in these figures therefore captures both the ACA policy variation, as well as any variation in penalties due to systematic differences in matching. Figure 2.3 shows the overall marriage penalties or subsidies by each partner's income for our entire sample. While the delineations between penalty and subsidy levels are less stark than what we see in the appendix figures, figure 2.3 reveals the same basic pattern: lower-income couples face lower penalties (or subsidies), while higher earners face

larger penalties; and the magnitude of the penalty varies both with individual and joint income of the partners. Figure 2.4 divides the sample by state's Medicaid expansion status. We observe a different pattern in the penalty distribution between couples in Medicaid expansion and non-expansion states. Couples in Medicaid non-expansion states are more likely to experience subsidies than similar couples in Medicaid expansion states. This is because low-earners in non-expansions states fall in the coverage gap, where they are eligible for neither Medicaid or premium tax credits, and thus face the full premium if they are unmarried. Marriage increases the ratio of income to FPL and thus renders them eligible for premium tax credits. Figure 2.5 shows the marriage penalties by number of children. The general pattern of Figure 2.5 shows that, among similar income combinations, marriage penalties are smaller when partners have a greater number of children. This is because getting married with more children does not increase the ratio of income to the FPL much, since FPL is larger when there are more children. Figure 2.6 shows the marriage penalties or subsidies by the age group. Older individuals face higher insurance premiums, and thus have larger marriage penalties or subsidies than younger individuals.

2.4.3 Empirical model

We use the penalties we computed for each couple in a difference-in-differences framework. Our identification strategy compares the marital status of individuals facing different amounts of ACA marriage penalties, before and after the ACA policy became effective. For example, we compare the marital status of a couple facing a $1,000 marriage penalty after 2014 to the marital

status of other couples with larger and smaller penalties post-2014, and also to the marital status of a similar couple pre-2014 who would have faced a $1,000 penalty had the ACA been in place.

We use the following model to estimate the policy impacts on marital status:

$$Y_i = \beta_0 + \beta_1(Penalty_i * Post_t) + \beta_2 Penalty_i + \beta_3 X_i + \lambda_s + \gamma_t + \theta_a + \mu_k + \varepsilon_i \quad (2.1)$$

Where Y_i is a dummy variable that equals one if a couple *i* living in state *s* during year *t* reported being married, and zero otherwise. $Penalty_i$ is a continuous variable that represents the amount of marriage penalty or subsidy that the couple faces. $Post_t$ is a dummy variable that equals one if the observation comes from 2014 and later. The vector X_i includes couple characteristics such as own gender, race, education, income and partner's income. The model also includes state fixed effects λ_s to account for the time-invariant state-specific characteristics that affect marital status, year fixed effects γ_t that account for the time-variant effects that apply to all states, age fixed effects θ_a, and number of children fixed effects μ_k. The error term is presented by ε_i. Standard errors are clustered at state level to account for correlated observations. The coefficient β_1 is the key estimate that represents the policy effect on the marriage outcome. It captures the conditional change in marital status associated with a $1,000 increase in marriage penalty. We can interpret β_1 as an estimate of the causal effect of the ACA marriage penalty on marriage if we assume that in the absence of the ACA, trends in marital status across different couple-types facing different levels of penalties would have remained constant.

If, however, marriage penalties or subsidies are correlated with changes in characteristics that also affect the marriage decision, then the estimate of β_1 could be biased. This would occur if there are trends in marriage rates particular to couples who have systematically high or low penalties. For example, if both partners work and earn similar high incomes, they will face a

higher marriage penalty. If trends in marital status also differ by income level, then we might spuriously conclude that higher marriage penalties are associated with marriage. Furthermore, if couples change characteristics in response to the ACA, then our estimate will capture these behavioral responses in addition to any actual relationship between the penalties and marriage. Most worrisome is the case where couples change their income in response to the ACA. Because the ACA uncoupled insurance coverage from work for many Americans, this is a likely scenario.

To address this potential endogeneity issue, we use a two-stage simulated instrumental variables (IV) approach. The idea here is to create a measure of policy-induced variation in ACA marriage penalties that is independent of individual couples' income. To do this, we take a random sample of 100,000 couples from the 2011 sample. Next, using the actual earnings that we observe, we run a simulation that identifies what each couple's marriage penalty would have been if they had lived in each potential combination of state and year, number of children and age group. This creates a set of 7140 hypothetical marriage penalties for each couple (51 states X 7 years X 4 fertility groups X 5 age groups). We average the set of hypothetical penalties at the state, year, number of children, and age group level, creating the simulated instrument ($Exposure_{stka}$). Unlike the $Penalty_i$ variable, which is determined in part by partner- or match-specific attributes of each couple at the time we observe them, $Exposure_{stka}$ is independent of these factors since it is calculated using the same sample with fixed characteristics. We merge $Exposure_{stka}$ into our full dataset by state, year, number of children and age group. Thus, we now have a measure of the policy-induced variation in marriage penalties for each couple in the full dataset. Figure 2.2 shows the distribution of $Exposure_{stka}$. Note, $Exposure_{stka}$ is

independent of income, and thus has less variation than the actual marriage penalties. Using $Exposure_{stka}$, we estimate the following two-stage instrumental variable model:

First stages:

$$Penalty_i = \alpha_0 + \alpha_1 Exposure_{stka} * Post_t + \alpha_2 Exposure_{stka} + \alpha_3 X_i + \lambda_s + \gamma_t + \theta_a + \mu_k + \varepsilon_i \quad (2.2)$$

$$Penalty_i * Post_t = \delta_0 + \delta_1 Exposure_{skta} * Post_t + \delta_2 Exposure_{skta} + \delta_3 X_i + \lambda_s + \gamma_t + \theta_a + \mu_k + \varepsilon_i \quad (2.3)$$

Second stage:

$$Y_i = \beta_0 + \beta_1 (\widehat{Penalty_i} * Post_t) + \beta_2 \widehat{Penalty_i} + \beta_3 X_i + \lambda_s + \gamma_t + \theta_a + \mu_k + \varepsilon_i \quad (2.4)$$

In equation 2.4, β_1 again provides an estimate of the effect of a $1,000 increase in marriage penalties on the marriage rate. However, because the variation in $\widehat{Penalty_i}$ is driven by the policy-induced variation in penalties, we need not be as concerned that it is biased by behavioral responses to the ACA in earnings or other factors.

2.5 Results

2.5.1 Main results

Figure 2.7 plots the average marriage rate by the ACA marriage penalty exposure $Exposure_{stka}$ for couples whom we observe before and after 2014. Negative value means subsidies and positive value means penalties. Exposure to the penalty in the pre-ACA years is hypothetical, representing the exposure that couples would have faced had the ACA been effective. This group acts as a control group, demonstrating any underlying correlation between policy exposure and

marriage rates driven by different baseline marriage rates among couples in different states, with different numbers of children or of different ages. As can be seen from the light grey line in the graph, there is positive association between penalty magnitude and marriage rates in the pre-2014 era. The black line plots the relationship between policy exposure to penalties and the marriage rate in the policy-effective period. If couples respond to the penalties in their marriage decisions, we expect to see higher marriage rates for couples with negative penalties (subsidies) in the post-2014 era, and lower marriage rates for couples with positive penalties. The figure confirms this hypothesis: when penalty exposure is positive, the percentage of married couples is consistently lower in post-2014 years than in the pre-2014 years. When penalty exposure is negative (subsidies), the percentage of married couples is slightly higher than in post-2014 years than in the pre-2014 era. When the penalty is close to zero, the probabilities of being married in two time periods are similar.

Table 2.2 demonstrates this result in regression results. We present the estimated effects of the ACA marriage penalties or subsidies on marital status for the full sample, and then separately by gender. The dependent variable is a dummy variable that equals 1 if the couple is married, and 0 otherwise. The coefficient on the interaction $Penalty_i * Post_t$ is the key coefficient of interest, which represents the association between the ACA marriage penalty and marital status. The coefficients represent the estimated effect of a $1,000 increase in marriage penalties. All models include a set of demographic controls, such as race/ethnicity, education, own income, and partner's income. Models also include state fixed effects, year fixed effects, age fixed effects, and number of children fixed effects.

Columns 1 through 3 show the relationship between the ACA marriage penalty and marital status using OLS, ignoring the potential endogeneity issue. These results shows that, for the full sample, the probability of being married drops by 0.7 percentage points with a $1,000 increase in the marriage penalty – about a 1 percent reduction in the marriage rate on a baseline rate of 64 percent. We estimate a larger coefficient for men than for women. We find that a $1,000 increase in penalty is associated with a 1 percent decrease in the marriage rate for men (0.69 percentage points on a baseline marriage rate of 67 percent); and a 0.6 percent reduction for women (0.36 percentage points on a baseline marriage rate of 61 percent).

Columns 4 through 6 shows the estimated effects using the IV approach. When we use the simulated penalty exposure to instrument for the actual marriage penalty, the estimated effects become larger: the probability of being married drops by about 5.9 percentage points with a $1,000 increase in marriage penalty reduced. With an average marriage rate of 64 percent, this represents an 9.2 percent decrease in the marriage rate. Furthermore, when we use the IV approach, the effects for men and women become nearly identical in relative magnitude.

The above results show the estimated impact of the ACA marriage penalty on marital status, under the assumption that children are assigned to women when married couples divorce. We conduct a sensitivity check to verify whether the estimated effects are robust to using different assignments of children. Table 2.3 represents the results of IV approach. Columns 1 though 3 show the estimated effect of the ACA marriage penalty on the probability of being married under the assumption that children are assigned to men when we assume married couples are single. The estimated effect is still significant and negative, but smaller in magnitude than when we assign children to women. For the full sample, the probability of being married

drops by 4.68 percentage points with a $1,000 increase in marriage penalty. The effect is slightly larger for men (5.34 percentage points) than women (3.85 percentage points). Columns 4 through 6 show the results under the assumption that children are assigned to the higher earner when we assume married couples are single. Because men tend to be the higher earners, these estimates are very similar to those in columns 1 through 3.[41] Overall, Table 2.3 shows that our result is not highly sensitive to the marriage penalty calculation method.

2.5.2 Heterogenous effect

We check whether the policy had differential effects among different groups by running separate regressions for each group. Marriage rates vary widely by demographic characteristics such as education and race/ethnicity (Marsh and Woods 2019). The divergence in marriage can be linked to the differences in marriage desire and social and economic barriers to marriage. For example, Burdette et al (2011) shows that marital intention is significantly higher among people with a college degree and above than less educated people. Raley et al (2016) shows Black people and Hispanic women have less incentive to wed than White people. Kuo and Raley (2016) points out that racial/ethnic minorities are more worried about the quality of marital relationship because of many factors, including drug abuse, domestic violence, and employment prospects. In addition, the marriage decision of less educated people and racial/ethnic minorities normally are more sensitive to financial reasons including receiving social welfare (Parker and Stepler, 2017; Eissa and Hoynes 2000). For these reasons, the ACA marriage penalty/subsidy may have heterogenous effect on the marital status across education and racial/ethnic groups.

[41] In the sample, more than 80 percent of married men earn more than married women.

Table 2.4 shows results by educational attainment. Columns 1 through 3 show the IV estimates for individuals with a college degree, 28 percent of our sample. Columns 4 through 6 show the estimates for individuals without a college degree. As can be seen, the ACA marriage penalty has larger effects on the group with less education. A $1,000 increase in marriage penalty reduced the probability of being married by 7.34 percentage points among the no-college group, and by 2.96 percentage points among the group with college degree. With the average marriage rate of 75 percent among the group with college degree and 59 percent among the no-college group, the above estimates represent 12.4 percent and 3.9 percent decreases in marriage rates, respective. Among individuals with a college degree, women are more affected by the penalty, with a 3.43 percentage points (4.8 percent) decline in the marriage rate. Among individuals without college degree, men are more likely to be affected, with a 7.93 percentage points (12.8 percent) decline in the marriage rate.

Table 2.5 shows results by race/ethnicity. Columns 1 through 3 show the IV estimates for people who identify as White people, and columns 4 through 6 show the estimates for those who identify as a racial category other than White. The ACA marriage penalty has larger impact on racial/ethnic category other than White people. A $1,000 increase in marriage penalty reduced the probability of being married by 8.7 percentage points (17 percent) among racial/ethnic groups other than White, and by 4.83 percentage points (7.2 percent) among White people. Among people identifying as White, the marriage rate response to the ACA marriage penalty is similar to men and women (7.2 percent decline for men and 6.9 percent decline for women). However, the marriage rate among men in racial/ethnic groups other than White is more

responsive than the marriage rate among women in racial/ethnic groups other than White (18.2 percent decline for men and 15.8 percent decline for women).

2.6 Conclusion

This paper examines the effect of the ACA Medicaid expansion and premium tax credits on marital status. We illustrate that people face financial penalties or subsidies when getting married as a result of the ACA policy structure. Our empirical evidence shows that the probability of being married drops by about 9 percent with a $1,000 increase in marriage penalty. The effect is larger among less educated group and racial/ethnic minorities. Men and women in particular groups responded differently, as well. For example, the marriage rate of highly educated women was more affected than highly educated men, while it was the opposite in the less educated group.

We also find that our IV analysis, which accounts for the fact that individuals might adjust income in response to the ACA, produces significantly larger estimates than the OLS approach. The downward bias of the OLS estimates imply that couples with larger marriage penalties also increased their marriage rates after ACA. This suggests that higher earning couples – who tend to have higher baseline marriage rates than lower-earning couples – were better able to shift labor force decisions in response to the ACA.

The marriage rate has been declining for decades in the U.S., with larger reductions among less educated people (Parker and Stepler, 2017). Couples in this group choose non-marital cohabitation. This has raised concerns – for example, increasing births to unmarried women (Popenoe, 2009). There is a long line of literature linking marriage with better health, more wealth, and better child outcomes (e.g., Waite and Gallagher, 2000). Governments in the

U.S. have been devoted to promoting marriage, including efforts in tax reform and other family supports (Fisher, 2003). Our study shows that the recent largest health insurance expansion, which aims at increasing health insurance coverage, may further deter marriage. We also demonstrate that different types of couples are more or less affected by the policy. However, our study does not imply whether the negative effect of the ACA health insurance expansion on marriage is good or bad and whether the design of the programs should be changed.

This study only focuses on the marital status at the time of the survey. Future research efforts can be made to disentangle the sources that contributed to the lower marriage rate, for example, increases in never married choosing not to marry, versus more people divorcing, versus unmarried cohabitating couples delaying marriage.

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Table 2.1. Descriptive summary

	Married	Cohabiting	Single
Men			
Age	47.27	41.71	45.47
White	0.75	0.67	0.64
College	0.32	0.18	0.17
Earnings (2017$)	57584.07	36575.59	20881.16
Women			
Age	45.30	39.80	46.34
White	0.74	0.68	0.57
College	0.34	0.24	0.20
Earnings (2017$)	26451.52	26965.13	21519.70
Household			
Earnings (2017$)	84314.54	63540.47	54421.31
No. of dependent children	0.97	0.39	0.34
Marriage penalty (2017$)	1679.80	482.71	784.24
Marriage penalty exposure (2017$)	1318.78	844.43	879.13
Proportion of facing marriage penalty	0.69	0.58	0.46
Obs	4320363	457591	2000405

Note: ACS 2011-2017; Age 26-64; Individual income is below 400% of FPL (assign children to women if divorced); combined income < $250,000;

Table 2.2. Effect of the ACA marriage penalty on the probability of being married

	OLS			IV		
	Full	Men	Women	Full	Men	Women
$Penalty_{ist} * Post_t$	**-0.0070*** (0.0008)**	**-0.0069*** (0.0009)**	**-0.0036*** (0.0008)**	**-0.0582*** (0.0086)**	**-0.0616*** (0.0094)**	**-0.0527*** (0.0077)**
Male	0.0115*** (0.0030)			0.0100*** (0.0034)		
Race/ethnicity (reference: White)						
Black	-0.2476*** (0.0082)	-0.1586*** (0.0080)	-0.2966*** (0.0081)	-0.2247*** (0.0091)	-0.1265*** (0.0092)	-0.2885*** (0.0088)
Hispanic	-0.0196 (0.0120)	0.0072 (0.0102)	-0.0472*** (0.0135)	-0.0087 (0.0151)	0.0205 (0.0145)	-0.0363** (0.0162)
Asian/other	0.0129 (0.0118)	0.0139 (0.0130)	0.0029 (0.0111)	0.0246* (0.0131)	0.0265* (0.0150)	0.0185 (0.0120)
Have college degree	0.0468*** (0.0023)	0.0326*** (0.0024)	0.0614*** (0.0025)	0.0402*** (0.0036)	0.0093** (0.0040)	0.0623*** (0.0039)
Own income ($1000)	0.0033*** (0.0001)	0.0039*** (0.0001)	0.0016*** (0.0001)	0.0002 (0.0003)	0.0001 (0.0003)	-0.0001 (0.0003)
Partner's income ($1000)	0.0018*** (0.0001)	-0.0002* (0.0001)	0.0027*** (0.0001)	-0.0023*** (0.0003)	-0.0039*** (0.0004)	-0.0015*** (0.0003)
Number of children (reference: none)						
One child	0.0849*** (0.0049)	0.1017*** (0.0041)	0.0624*** (0.0053)	0.0321*** (0.0104)	0.0611*** (0.0097)	0.0032 (0.0106)
Two children	0.1753*** (0.0057)	0.1783*** (0.0045)	0.1557*** (0.0064)	0.0839*** (0.0143)	0.0987*** (0.0140)	0.0637*** (0.0142)
Three and more children	0.1827*** (0.0070)	0.1772*** (0.0062)	0.1597*** (0.0079)	0.0610*** (0.0189)	0.0705*** (0.0188)	0.0413** (0.0189)
FE(state, age, year)	yes	yes	yes	yes	yes	yes
First stage F-statistic > 10				yes	yes	yes
Mean of being married	0.64	0.67	0.61	0.64	0.67	0.61
Obs	6778359	3244816	3533543	6778359	3244816	3533543

Notes: Data from the ACS 2011-2017 are used to obtain these estimates. Dependent variable is marital status (dummy). Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 2.3. Sensitivity check (IV)

	Assign children to men			Assign children to higher earner		
	Full	Men	Women	Full	Men	Women
$Penalty_{ist} * Post_t$	**-0.0468****	**-0.0534****	**-0.0385****	**-0.0516****	**-0.0587****	**-0.0427****
	(0.0192)	**(0.0221)**	**(0.0153)**	**(0.0237)**	**(0.0272)**	**(0.0191)**
Male	0.0141***			0.0135***		
	(0.0032)			(0.0033)		
Race/ethnicity (reference: White)						
Black	-0.2515***	-0.1529***	-0.3013***	-0.2532***	-0.1548***	-0.3045***
	(0.0077)	(0.0078)	(0.0078)	(0.0077)	(0.0079)	(0.0078)
Hispanic	-0.0205	0.0076	-0.0492***	-0.0222*	0.0057	-0.0509***
	(0.0126)	(0.0108)	(0.0141)	(0.0124)	(0.0107)	(0.0140)
Asian/other	0.0170	0.0171	0.0066	0.0184	0.0184	0.0076
	(0.0121)	(0.0133)	(0.0116)	(0.0125)	(0.0137)	(0.0119)
Have college degree	0.0493***	0.0340***	0.0666***	0.0522***	0.0357***	0.0695***
	(0.0033)	(0.0036)	(0.0037)	(0.0039)	(0.0039)	(0.0043)
Own income ($1000)	0.0025***	0.0033***	0.0004	0.0025***	0.0031***	0.0004
	(0.0002)	(0.0001)	(0.0003)	(0.0002)	(0.0002)	(0.0003)
Partner's income ($1000)	0.0007***	-0.0018***	0.0016***	0.0006***	-0.0019***	0.0015***
	(0.0002)	(0.0003)	(0.0001)	(0.0002)	(0.0004)	(0.0001)
Number of children (reference: none)						
One child	0.0831***	0.1068***	0.0538***	0.0873***	0.1103***	0.0550***
	(0.0051)	(0.0042)	(0.0055)	(0.0047)	(0.0040)	(0.0050)
Two children	0.1803***	0.1889***	0.1512***	0.1911***	0.1993***	0.1582***
	(0.0057)	(0.0046)	(0.0064)	(0.0055)	(0.0046)	(0.0058)
Three and more children	0.1906***	0.1897***	0.1541***	0.2015***	0.2022***	0.1623***
	(0.0070)	(0.0063)	(0.0082)	(0.0068)	(0.0065)	(0.0076)
FE(state, age, year)	yes	yes	yes	yes	yes	yes
First stage F-statistic > 10	yes	yes	yes	yes	yes	yes
Mean of being married	0.63	0.66	0.60	0.63	0.66	0.60
Obs	6717223	3214248	3502975	6659967	3185620	3474347

Notes: Data from the ACS 2011-2017 are used to obtain these estimates. Dependent variable is marital status (dummy). Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 2.4. Heterogenous effect (IV): Education

	Have college degree			Without college degree		
	Full	Men	Women	Full	Men	Women
$Penalty_{ist} * Post_t$	**-0.0287*** (0.0063)**	**-0.0244*** (0.0058)**	**-0.0343*** (0.0070)**	**-0.0725*** (0.0112)**	**-0.0793*** (0.0125)**	**-0.0623*** (0.0095)**
Male	-0.0411*** (0.0027)			0.0327*** (0.0048)		
Race/ethnicity (reference: White)						
Black	-0.2182*** (0.0086)	-0.1019*** (0.0070)	-0.2841*** (0.0084)	-0.2240*** (0.0092)	-0.1217*** (0.0093)	-0.2841*** (0.0090)
Hispanic	-0.0430*** (0.0094)	-0.0202*** (0.0074)	-0.0600*** (0.0108)	-0.0005 (0.0152)	0.0267* (0.0144)	-0.0305* (0.0167)
Asian/other	0.0287*** (0.0072)	0.0231*** (0.0087)	0.0340*** (0.0063)	0.0194 (0.0181)	0.0300 (0.0201)	0.0079 (0.0162)
Own income ($1000)	0.0005*** (0.0002)	0.0004** (0.0002)	0.0013*** (0.0001)	0.0000 (0.0004)	0.0005 (0.0004)	-0.0019*** (0.0004)
Partner's income ($1000)	-0.0019*** (0.0002)	-0.0029*** (0.0002)	-0.0018*** (0.0003)	-0.0025*** (0.0004)	-0.0049*** (0.0005)	-0.0011*** (0.0003)
Number of children (reference: none)						
One child	0.1276*** (0.0095)	0.1215*** (0.0090)	0.1247*** (0.0101)	0.0035 (0.0106)	0.0517*** (0.0097)	-0.0417*** (0.0108)
Two children	0.1454*** (0.0133)	0.1228*** (0.0126)	0.1538*** (0.0140)	0.0784*** (0.0145)	0.1174*** (0.0138)	0.0310** (0.0147)
Three and more children	0.0867*** (0.0182)	0.0620*** (0.0171)	0.0954*** (0.0194)	0.0847*** (0.0187)	0.1167*** (0.0181)	0.0336* (0.0191)
FE(state, age, year)	yes	yes	yes	yes	yes	yes
First stage F-statistic > 10				yes	yes	yes
Mean of being married	0.75	0.79	0.72	0.59	0.62	0.57
Obs	1914977	882806	1032171	4863382	2362010	2501372

Notes: Data from the ACS 2011-2017 are used to obtain these estimates. Dependent variable is marital status (dummy). Standard errors in parentheses are clustered at the state level. *, **, and *** indicate $p < .1$, $p < .05$, and $p < .01$ respectively.

Table 2.5. Heterogenous effect (IV): Race/ethnicity

	Whites			Racial/ethnic groups other than White		
	Full	Men	Women	Full	Men	Women
$\text{Penalty}_{ist} * \text{Post}_t$	**-0.0483*** (0.0066)**	**-0.0495*** (0.0070)**	**-0.0448*** (0.0061)**	**-0.0870*** (0.0140)**	**-0.1015*** (0.0174)**	**-0.0730*** (0.0115)**
Male	-0.0115*** (0.0025)			0.0606*** (0.0069)		
Race/ethnicity (reference: Asian/other)						
Black				-0.2367*** (0.0140)	-0.1245*** (0.0181)	-0.3110*** (0.0122)
Hispanic				-0.0445** (0.0187)	-0.0125 (0.0158)	-0.0681*** (0.0207)
Have college degree	0.0331*** (0.0031)	0.0085*** (0.0030)	0.0536*** (0.0041)	0.0617*** (0.0065)	0.0141* (0.0085)	0.0884*** (0.0060)
Own income ($1000)	0.0007*** (0.0003)	0.0008*** (0.0003)	0.0002 (0.0003)	-0.0012*** (0.0004)	-0.0020*** (0.0005)	-0.0008* (0.0004)
Partner's income ($1000)	-0.0015*** (0.0003)	-0.0030*** (0.0003)	-0.0008*** (0.0003)	-0.0045*** (0.0005)	-0.0065*** (0.0006)	-0.0033*** (0.0004)
Number of children (reference: none)						
One child	0.0392*** (0.0086)	0.0599*** (0.0078)	0.0154* (0.0091)	0.0190 (0.0169)	0.0684*** (0.0167)	-0.0202 (0.0164)
Two children	0.0875*** (0.0117)	0.0923*** (0.0114)	0.0749*** (0.0121)	0.0841*** (0.0223)	0.1277*** (0.0229)	0.0466** (0.0213)
Three and more children	0.0553*** (0.0154)	0.0487*** (0.0153)	0.0467*** (0.0158)	0.0830*** (0.0294)	0.1372*** (0.0313)	0.0367 (0.0275)
FE(state, age, year)	yes	yes	yes	yes	yes	yes
First stage F-statistic > 10				yes	yes	yes
Mean of being married	0.67	0.69	0.66	0.51	0.56	0.46
Obs	4733745	2323316	2410429	2044614	921500	1123114

Notes: Data from the ACS 2011-2017 are used to obtain these estimates. Dependent variable is marital status (dummy). Standard errors in parentheses are clustered at the state level. *, **, and *** indicate p < .1, p < .05, and p < .01 respectively.

Figure 2.1. The distribution of $Penalty_i$

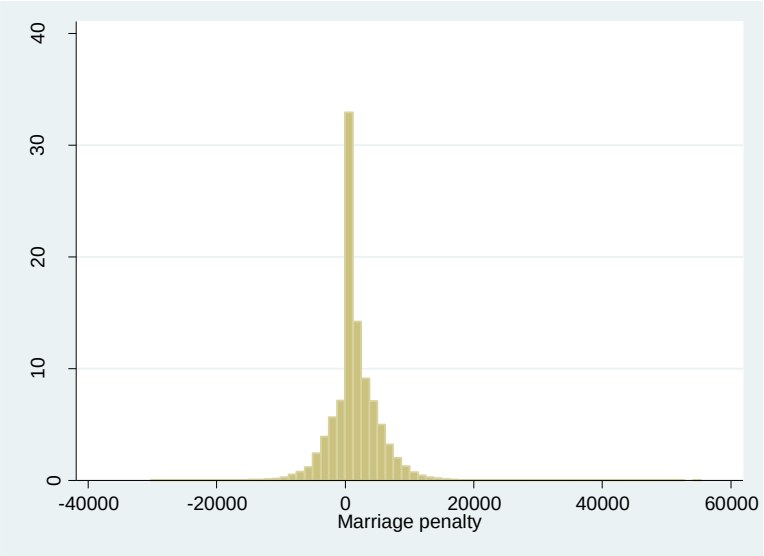


Figure 2.2. The distribution of $Exposure_{stka}$

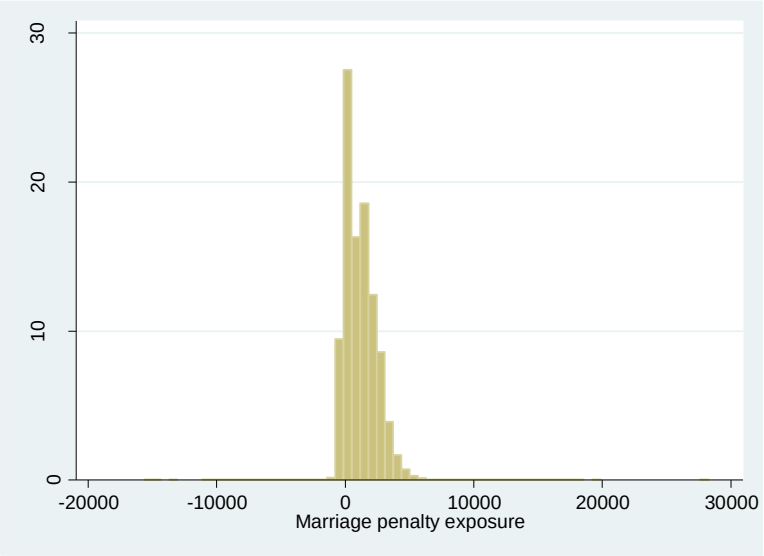


Figure 2.3. Overall marriage penalties or subsidies by each partner's income

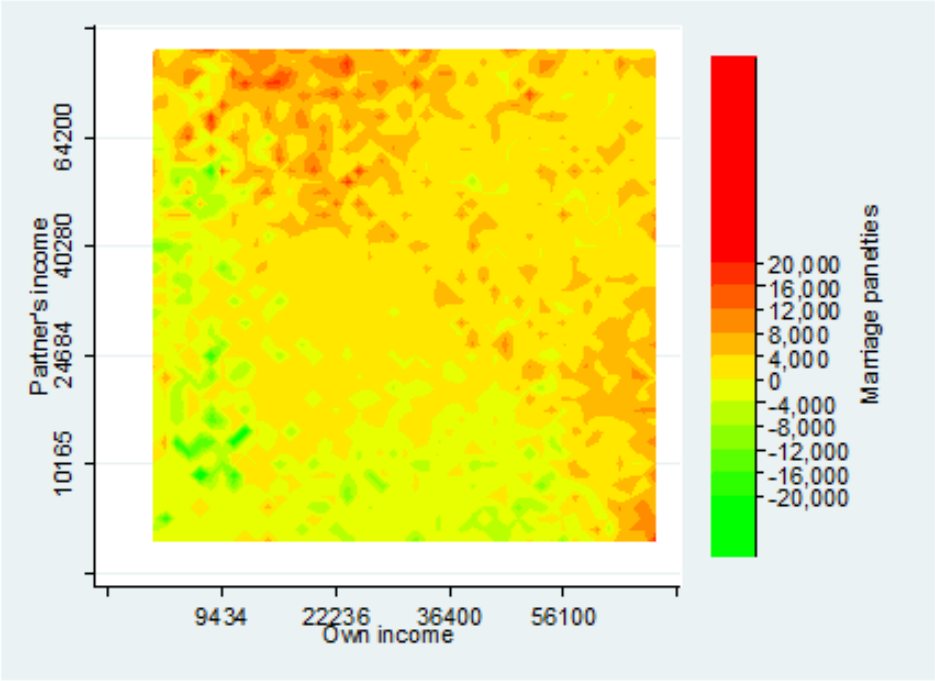


Figure 2.4. Marriage penalties or subsidies by each partner's income and state's Medicaid expansion status

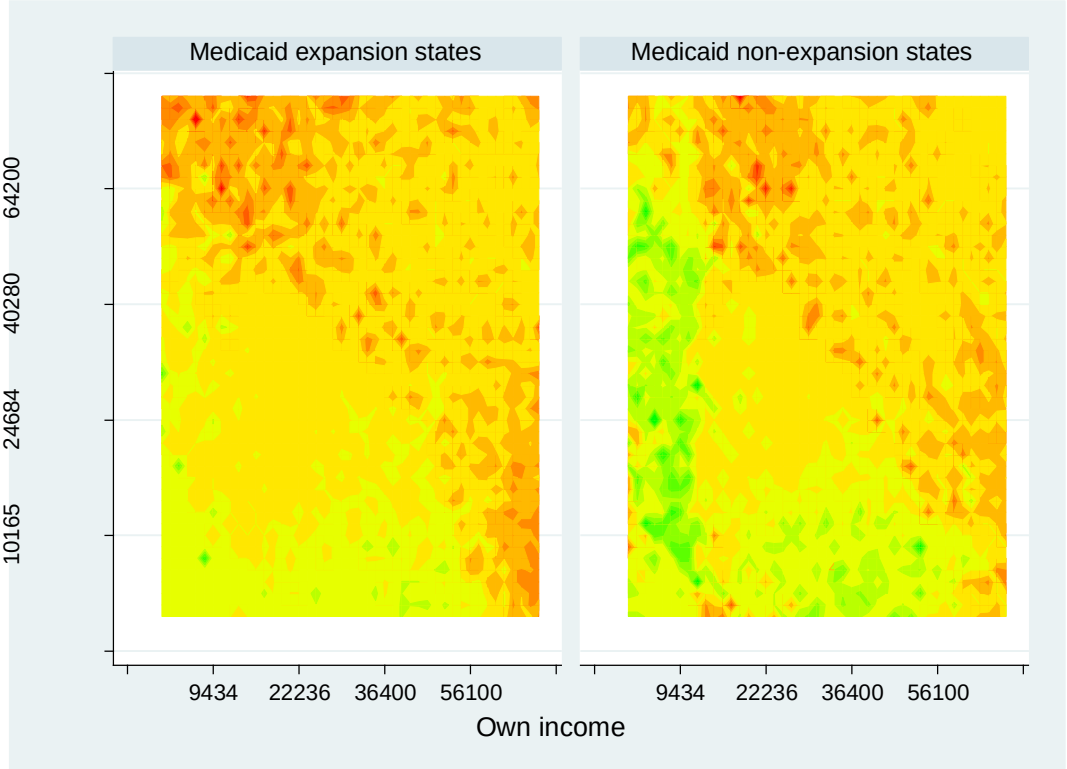


Figure 2.5. Marriage penalties or subsidies by each partner's income and number of children

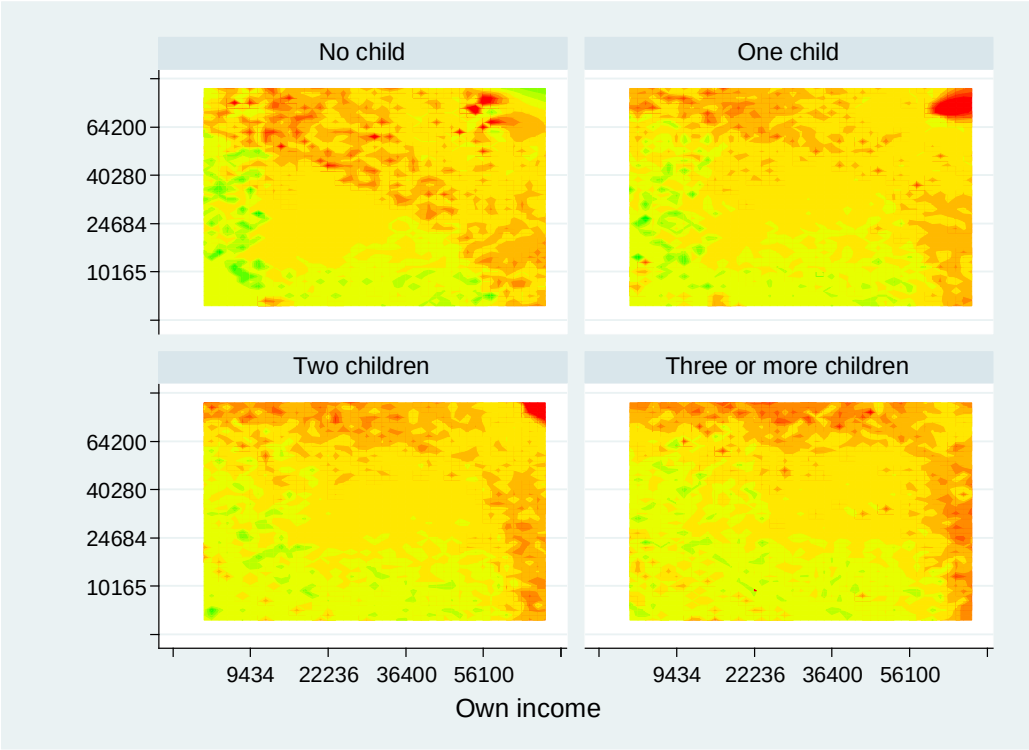


Figure 2.6. Marriage penalties or subsidies by each partner's income and age group

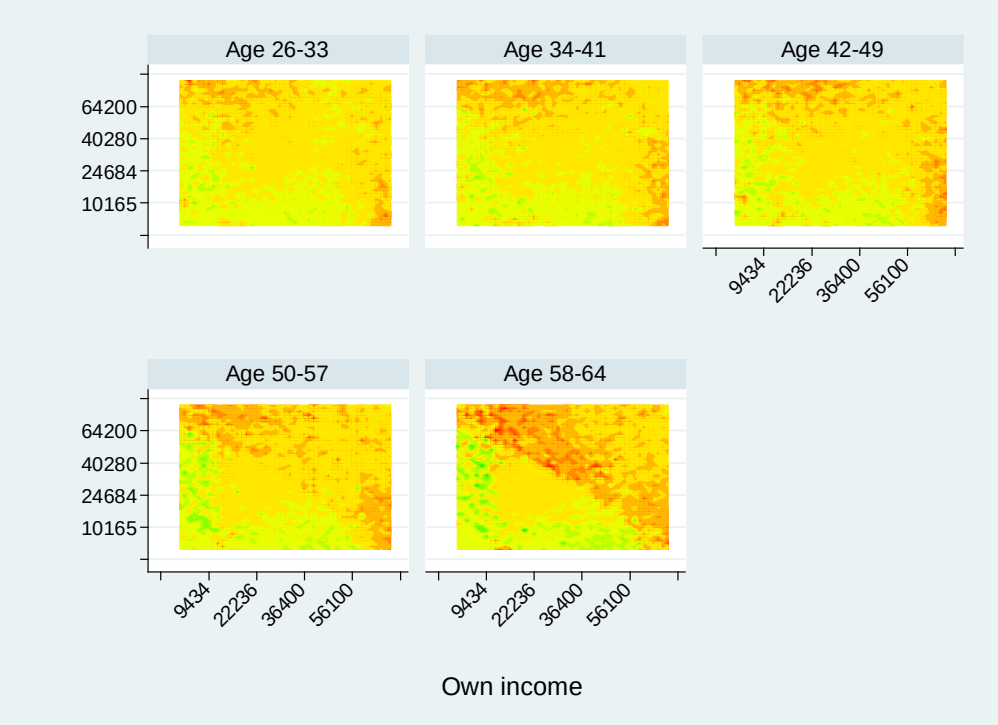
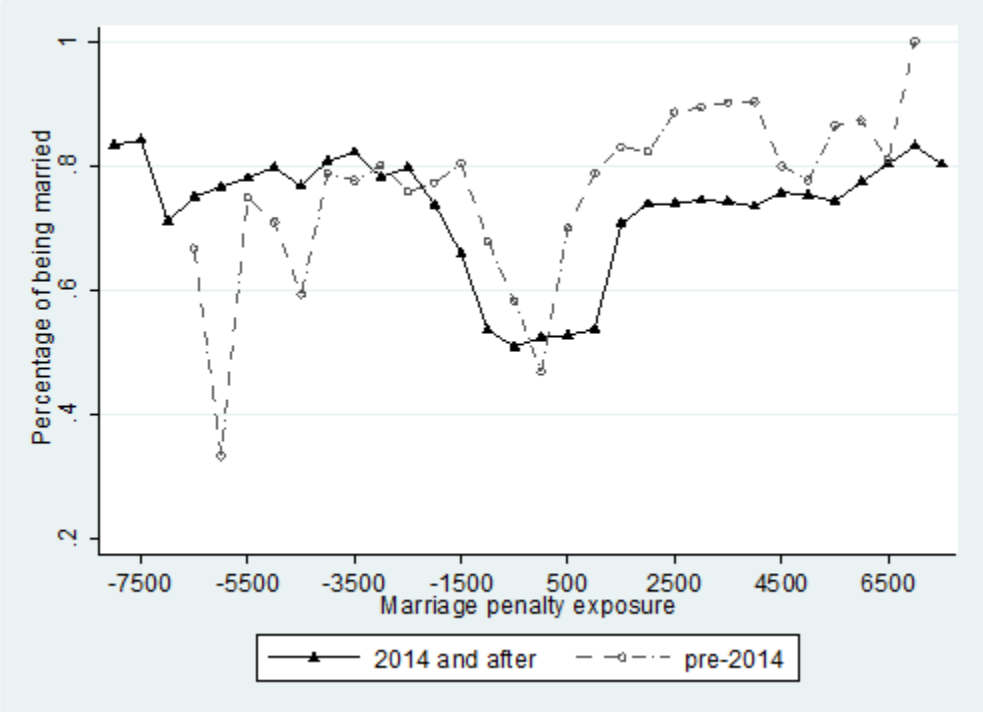


Figure 2.7. Trends of marital status by $Exposure_{stka}$



Chapter 3. Unemployed and Uninsured: Black-White disparities during the Covid-19 Recession

3.1 Introduction

The Covid-19 pandemic has caused unprecedented economic consequences. From March to May 2020, around 35 million workers were laid off. These layoffs are especially troublesome in the American context, where vast numbers of people have health insurance through their employers. Researchers estimate that 16.2 million have lost employer sponsored health insurance (ESI) in the U.S., and millions more will be added to the uninsured population as a result of future Covid-19 layoffs (Garrett and Gangopadhyaya, 2020a; Garrett and Gangopadhyaya, 2020b; Dorn, 2020;). Fortunately, the Affordable Care Act (ACA) provides a safety net that allows unemployed people and their families to regain health insurance coverage.[42] However, due to the means-tested eligibility rules and the idiosyncrasies of the program structure, there have been concerns that the ACA benefits do not provide an adequate health insurance safety net for the newly unemployed (Blumberg et al., 2020a; Blumberg et al., 2020b; Garfield et al., 2020; Straw et al., 2020). This study explores how access to government subsidized health insurance coverage varies across people that have lost their ESI due to a Covid-19 layoff, and whether that variation has worsened Black-White inequities in access to subsidize health insurance.

[42] From 2011-2013 to 2014-2018, uninsured rates dropped significantly among unemployed people (Gangopadhyaya and Garrett, 2020).

We explore this question from three perspectives. First, we explore the research question through a lens of racial/ethnic equity. Existing estimates suggest that people who identify with racial/ethnic groups other than White have been disproportionately affected by Covid-19 layoffs (Fairlie et al., 2020; Montenovo et al., 2020). Thus, underlying racial/ethnic inequities in healthcare access may potentially be worsened by Covid-19 layoffs. In this study, we focus on the Black-White disparity in access to ACA benefits after losing a job for two reasons. First, ACA eligibility is restricted by immigration status. Undocumented immigrants are not eligible for any ACA benefits, while noncitizen lawful immigrants can be eligible if they have lived in the United States for more than five years.[43] Census surveys, however,do not include information on immigration status, making it difficult to precisely estimate a respondent's ACA eligibility. We thus only focus on Black and White people, since a large percentage of Hispanic and Asian people are immigrants.[44] Second, there is an especially worrisome possibility given that Black people are more likely to contract and die of Covid-19 (e.g., Yancy, 2020; Abedi et al., 2020). Black and White people have, on average, different household structures, different income distributions, and live in different places, access to government assistance with health insurance after a Covid-19 layoff and loss of ESI will therefore differ systematically between Black and White people. Using this lens, we analyze whether Covid-19 layoffs interact with gaps in the safety net to narrow or widen racial/ethnic disparities in healthcare access. Second, we evaluate the extent to which Covid-19 federal aid has helped or hindered healthcare access for people who lost ESI due to a Covid-19 layoff. The Coronavirus Aid, Relief, and Economic Security (CARES) Act did not directly address how

[43] For more details, please see https://www.healthcare.gov/immigrants/lawfully-present-immigrants/
[44] In 2017, about 33 percent of Hispanics are immigrants (Noe-Bustamante, 2019), 59 percent of Asians are foreign born (Lopez et al., 2017), while 9 percent Black people are foreign born (Anderson and Lopez, 2018).

people who lost ESI could regain health insurance. However, under the CARES Act, unemployed people received the Federal Pandemic Unemployment Compensation (FPUC) [45] - an additional $600 per week cash on top of regular state unemployment benefits from April to July 2020, with the benefit period extended by 13 weeks. By increasing income, the FPUC program may have inadvertently affected access to ACA benefits – increasing access for some people, while jeopardizing access for others. Due to the fact that many states did not expand Medicaid, some people who lost ESI in these states would not have been eligible for Medicaid, or for premium subsidy – falling into the so-called "coverage gap". For these people, additional FPUC income may have moved them out of the coverage gap, increasing their eligibility for a premium subsidy. On the other hand, additional FPUC income may have reduced the benefits available to other people – by reducing their eligible premium subsidy, causing them to earn-out of the cost-sharing range of the premium subsidy program, or causing them to earn-out of the entire subsidy range by increasing their income beyond 400 percent of the federal poverty line (FPL). With a focus on Black-White differences, we explore how the FPUC income affected access to ACA benefits for Black and White people who lost ESI due to Covid-19 layoffs, as well as how the FPUC narrows or widens Black-White disparity in health insurance coverage.

Finally, given the uncertainties in the future labor market and continuity of government unemployment compensation, we estimate the access to ACA benefits for people who lost job and ESI in Covid-19, in the worst case, remain unemployed in 2021 and exhaust all unemployment compensation. Again, we explore this question through a lens of Black-White equity. We estimate how the worse scenario will affect Black-White disparity in access to

[45] CARES Act creates Pandemic Emergency Unemployment Compensation (PEUC) program and Pandemic Unemployment Assistance (PUA) program to assist workers who lost jobs. For more information, see https://www.dol.gov/coronavirus/unemployment-insurance.

government subsidized health insurance. Answering these questions has important implications for creating solutions to provide a more adequate safety net that addresses the Black-White disparity in health insurance coverage, especially during recessions.

To address these three objectives, we conduct a simulation-based analysis. Data limitations present a vital roadblock to obtaining up-to-the-minute estimates of relevant Covid-19 parameters. To overcome this limitation, we use data from both the 2018 American Community Survey (ACS) and the May 2020 Current Population Survey (CPS) and conduct an analysis in three stages. First, we use data from the May 2020 CPS to estimate the determinants of being laid off in the first moths of Covid-19. Next, we use the parameters obtained from the 2020 May CPS and apply them to the 2018 ACS data to identify those who are likely to be laid off in April and May 2020. Since the 2018 ACS data include detailed information on income and health insurance status – information missing from the 2020 May CPS – we can use these data to identify a sample of people who are likely to lose their ESI when Covid-19 hits. Finally, using our job-loss predications, we use the 2018 ACS data to estimate the ACA eligibility for those likely to be laid off based on their income before and after job loss, state of residence, and family status. We estimate the ACA eligibility based on three scenarios. First, we evaluate the ACA eligibility in 2020 if the FPUC had not provided any income. Second, we access the ACA eligibility in 2020 if newly unemployed received the FPUC. Finally, we estimate the ACA eligibility in 2021 if both state and federal unemployment compensation are exhausted. Our analysis focuses on identifying those who are ACA eligible because of being eligible for Medicaid, or being eligible for a premium subsidy with cost-sharing or without cost-sharing, and those who are ACA ineligible because of falling in the coverage gap or having family income beyond the premium subsidy cut-off at 400 percent of the FPL. All analysis is framed through

the lens of Black-White equity, emphasizing the extent to how various scenarios affect Black-White disparity in health insurance coverage.

Several existing studies explore how the FPUC affects ACA eligibility for workers who lost jobs and ESIs (e.g., Garfield et al., 2020; Blumberg et al., 2020a; Blumberg et al., 2020b). However, these studies have not explored racial/ethnic disparity in ACA eligibility under Covid-19. Racial/ethnic disparities in health insurance coverage and use of healthcare in the U.S. are well-known (Lillie-Blanton and Hoffman, 2005; Buchmueller et al., 2016; Chen et al., 2016). While the ACA's coverage expansions have led to an increase in coverage and access to healthcare, significant gaps between Black and White people remain (Baumgartner et al., 2020; Buchmueller and Levy 2020). Our analysis adds to the emerging knowledge on how the Covid-19 pandemic and government aid policies, interact with the existing ACA program rules by explicitly addressing issues of racial/ethnic equity.

Our results show that Black people who lost ESI would be more likely to receive no government assistance in health insurance than White people under Covid-19 if the FPUC had not provided any income. The main reason is that Black people who lost ESI would be more likely than White people to fall into the coverage gap. This is likely to widen the existing Black-White disparity in health insurance coverage, as people in coverage gap are low-income and can barely afford non-group private health insurance. The $600 per week FPUC that ended in July reduced the potential increase in the Black-White disparity in health insurance coverage mentioned above by increasing the access to premium subsidies for the majority of people who would fall in the coverage gap. Finally, in the worst case that newly unemployed exhaust all unemployment compensation in 2021, we find that the Black-White disparity in health insurance

coverage will be substantially exacerbated if no relevant government action is taken. Almost all people who are likely to be uninsured are those at-risk of falling in the coverage gap.

Given the uncertainties of future economic recovery and continuity of government unemployment compensation, we articulate that improving ACA provisions to provide an adequate safety net is particularly essential regarding the Black-White disparity in health insurance coverage. Solutions can include expanding Medicaid in the non-expansions states or expanding the income range for premium subsidy. Either way will allow more Black people at-risk of falling into the coverage gap to have access to more affordable health insurance.

3.2 Background

Federal legislation to remediate the economic consequences of Covid-19 pandemic so far have not addressed the issue of improving health insurance coverage among the newly unemployed. President Trump announced a program that reimburses healthcare providers for testing and treating uninsured people with Covid-19 using fund from Covid-19 relief package under the Coronavirus Aid, Relief, and Economic Security (CARES) Act, but it has been widely reported at the program falls short of promise.[46]

Workers losing employer sponsored health insurance (ESI) during Covid-19 pandemic may be eligible for government assistance with health insurance facilitated by the Affordable Care Act (ACA), including Medicaid and premium subsidy. The main factors affecting eligibility for ACA programs include family income, state of residence, and family status. The CARES Act provided an additional $600 per week in unemployment compensation (FPUC) on top of regular

[46] See https://www.nytimes.com/2020/08/29/health/Covid-obamacare-uninsured.html

state unemployment benefits for up to 17 weeks (from April to July). Regular state unemployment benefits have been extended from maximum 26 weeks to 39 weeks in most states. The extra unemployment compensation counts as income in determining some aspects of ACA eligibility.

Medicaid is a public health insurance with zero premiums and low cost-sharing. Medicaid expansion is one major provision under the ACA, which increased income eligibility cut-off to 138 percent of the FPL. However, not all states adopted Medicaid expansion. People in Medicaid expansion states with income below 138 percent of the FPL are all eligible for this public insurance. Few people in Medicaid non-expansion states are eligible because income eligibility cut-offs in those states are very low and childless adults are generally ineligible regardless of income. People losing ESI during Covid-19 may get Medicaid, especially those who live in Medicaid expansion states. Eligibility is evaluated monthly based on current family income, including state unemployment benefits for the unemployed person, plus any additional earnings from other family members who are still employed. The $600 per week FPUC and income earned in the calendar year before the job loss do not affect Medicaid eligibility.

Some other workers who lost ESI and who are not eligible for Medicaid may be eligible for premium subsidy. Provided by the federal government to low-to-moderate income people (income between 100-400 percent of the FPL) who do not have affordable ESI, the subsidies reduce the cost of purchasing private health insurance through the health insurance Marketplaces. People eligible for premium subsidy will contribute a certain percentage of their family income, or net premium. The contribution percentage increases from about 2 percent to

9.5 percent of income as income increases from between 100 and 400 percent of FPL.[47] People with income below 250 percent of the FPL are additionally eligible for cost-sharing reductions, which lowers out-of-pocket medical expenses like co-payments and deductibles.[48] Unlike Medicaid, eligibility for premium subsidies is based on projected annual income for the calendar year, which includes earnings before a job loss, and additional family income after job (unemployment benefits plus income from other family members). Note that the $600 per week FPUC is included in income when determining premium subsidy eligibility.[49]

While the ACA makes health insurance more affordable for many people, it does not reach everyone and thus leaves some people who have lost ESI during Covid-19 vulnerable. Among all people who have lost ESI, some may be eligible for neither Medicaid nor premium subsidy. These people include those who fall into the so-called coverage gap, or those who have annual household income above 400 percent of the FPL. The coverage gap includes a group of low-income people in Medicaid non-expansion states. Because the income eligibility cut-off for Medicaid is very low but income eligibility threshold for premium subsidy starts at 100 percent of the FPL, people with income in the gap in those states are not eligible for any government assistance with health insurance, and they are much likely to be uninsured. People with income above 400 percent of the FPL and thus not eligible for any financial assistance are also less likely to be insured, since premiums for these families will be large.

[47] The percentage may fluctuate a little bit year by year. In 2020, the lowest contribution percentage starts at 2.08 percent and the maximum contribution percentage is 9.78 percent.

[48] Without cost-sharing reduction, the maximum out-of-pocket medical spending is limit to $8,150 for an individual and $16,300 for a family. With cost-sharing reduction, the maximum out-of-pocket medical expending cannot exceed $2,700 for an individual and $5,400 for a family when income is below 200 percent of the FPL, and $6,500 for an individual and $13,000 for a family when income is between 200-250 percent of the FPL.

[49] Premium subsidies are advanceable, but it has to be reconciled based on actual income when people file tax return in the following year. People will either get refund or pay the difference between the actual subsidies they are eligible for and the subsidies they already received in the previous year. Cost-sharing reductions do not have to be reconciled when file tax return.

Given the huge impact of Covid-19 on employment and health, it is important to understand how people who have lost ESI because of job loss can access alternative health insurance coverage. A number of recent studies use simulation approaches to explore ACA eligibility or to predict health insurance coverage change as a result of Covid-19 layoff. Both parameters have important policy implication for improving health insurance coverage during the pandemic.

Blumberg et al. (2020a) estimates the ACA eligibility for workers and their dependents in vulnerable industries with high Covid-19 related unemployment. The study assumes all of these workers become unemployed from April to the end of 2020 and receive the maximum state unemployment benefits and FPUC. The authors conclude that receiving FPUC helps some vulnerable workers who originally might fall in the coverage gap become eligible for premium subsidy, but it also makes some vulnerable workers become ineligible for premium subsidy because FPUC makes their income too high to qualify.

Blumberg et al. (2020b) investigates the ACA eligibility for workers in vulnerable industries who had ESI before losing job. The study further extends the previous study by showing how additional FPUC proposed by the HEROES Act (provide $600 per week from April to December 2020 instead of ending in July), affects eligibility for premium subsidy. They also estimate the ACA eligibility when all unemployment compensation is exhausted (the year of 2021). They conclude that current FPUC helps more people who lost ESI in Medicaid non-expansion than those in Medicaid expansion states. Additional FPUC could make fewer people in both Medicaid expansion states and non-expansion states eligible for premium subsidy by increasing household annual income to above 400 percent of the FPL. At the beginning of 2021 when unemployed people exhaust all unemployment compensation and are still unemployed,

more than three quarters of those in Medicaid expansion states will be ACA eligible (mainly through Medicaid), while only half of those in Medicaid non-expansion states will be ACA eligible (mainly through premium subsidy).

Garfield et al. (2020) also explores the ACA eligibility for workers (and their families) who could potentially lose ESI. However, unlike the previous two studies, which assume all workers in vulnerable industries become unemployed, this study defines workers losing jobs using sampling probabilities based on the recent employment change by industry recorded by the 2020 CPS monthly data. In addition, instead of assuming all unemployed people receive the maximum state unemployment benefits, this study calculates the state unemployment benefits based on industry-specific distribution and state-specific generosity. Results show that current FPUC almost eliminated the coverage gap in Medicaid non-expansion states by helping recently unemployed with low income become eligible for premium subsidies. However, in 2021 when all unemployment compensation is exhausted, a substantial number of people in Medicaid non-expansion states will be in the coverage gap where they get no government assistance with health insurance, while the majority of people losing ESI under Covid-19 are Medicaid eligible in Medicaid expansions states.

Instead of estimating the ACA eligibility, some studies predict the health insurance coverage change during Covid-19 (Gangopadhyaya and Garrett 2020; Garrett and Gangopadhyaya, 2020; Dorn, 2020). Their basic idea is to simulate the health insurance coverage under different cases of unemployment (low, moderate, high) based on pre-pandemic health insurance coverage among unemployed workers. However, these studies do not account for the FPUC, which substantially affects the eligibility for premium subsidies.

To the best of our knowledge, none of the existing studies on health insurance coverage among newly unemployed during the Covid-19 recession explore racial/ethnic difference. Racial/ethnic disparities in health insurance coverage and use of healthcare are well-known. In addition, given the fact that Covid-19 has disproportionately affected the health of Black people, as both infection and death rates are much higher among Black people than White people (e.g., Tai et al., 2020; Millett et al., 2020), it is important to investigate how access to alternative health insurance coverage varies for Black versus White people who lost ESI.

The ACA assistance with health insurance coverage for recently unemployed Black and White people could be significantly different. Unemployed Black people are more likely than White people to fall in the coverage gap because a large proportion of them live in Medicaid non-expansion states. According to Artiga et al. (2020), the majority of Medicaid non-expansion states are located in the south, where the Black population is disproportionately high (greater than 15 percent). Unemployed White people are more likely to have higher annual household income than unemployed Black people, affecting their eligibility for ACA programs. Thus, Black and White people may have different barriers to access alternative health insurance when losing ESI during Covid-19. The FPUC, which increases income, can affect ACA eligibility for Black and White people who lost ESI in different ways.

Besides focusing on racial/ethnic differences, this study is also different from the relevant studies in two ways. First, instead of predicting workers' job loss based solely on the vulnerability of their industry, we use individual characteristics of newly unemployed people during Covid-19 to predict the probability of being laid off more precisely. It has been shown that younger and older workers, women workers, and less educated worker are more likely to be laid off (Montenovo et al., 2020). Second, we calculate the state unemployment benefits based

on worker's reported regular income before losing job, instead of assuming they receive the state maximum amount or using the industry-specific distribution.

3.3 Data and Method

We use data from the 2020 Basic Monthly Current Population Survey (CPS) in May to construct predict characteristics associated with job loss during Covid-19 pandemic. The basic Monthly CPS is a household-based survey conducted by the US Census Bureau that includes rich information about individual labor force supply. CPS asks unemployed individuals how many consecutive weeks that they have been without a job. We define an individual as a newly unemployed if he/she answered the number of consecutive weeks of being unemployed is not greater than 8 weeks. The Basic Monthly CPS data is generally collected in the calendar week that contains the 12th day of the month. The 2020 May CPS data is collected between May 10-16th. Thus, newly unemployed individuals represent those losing job between mid-March and mid-May 2020.

We limit the CPS sample to nonelderly adults in labor force. We use a logistic regression to measure the relationship between becoming newly unemployed and individual characteristics, including race/ethnicity, age, age square, gender, education, marital status, occupation, and industry. We also include an indicator of whether the individual worked in non-essential businesses that were mandated to be closed during Covid-19. Following Fairlie et al (2020), we define whether an industry is non-essential based on the criteria in the state of Delaware.[50] Table C.1 shows the regression estimates of our logistic model. Younger workers, women, racial/ethnic

[50] Delaware posted a comprehensive set of industry codes that can be matched to the same industry codes in the CPS, with comments about whether the industry is essential or not.

groups other than White (Black, Hispanic, and Asian/other), lower educated, unmarried, and workers in non-essential business are more likely to be laid off during Covid-19. The results are consistent with recent studies looking at Covid-19 job losses (e.g., Montenovo et al. 2020).

The May 2020 CPS data are well-suited to predicting who lost a job during the early months of Covid-19. However, the data lack detail on our question of interest: health insurance coverage. Thus, we use the coefficients derived from the logistic regression above on data from the 2018 American Community Survey. The ACS includes the covariates we use in our logistic regression above, as well as detailed information about health insurance status and income. We apply the coefficients from the CPS, and predict the probability of losing job under Covid-19 for individuals in 2018 ACS. Our sample of interest includes Black and White people who lost ESI because of job loss in the first months of Covid-19. To identify these people, we use the sample of nonelderly Black and White people in a family where at least one worker has a predicted layoff probability in the top 10 percentile of the predicted job loss distribution. We keep people that were previously covered by ESI and that would not be able to retain ESI through another family member after a job loss occurred in the family.[51] That leads to a final sample size of 54,351 White people and 13,374 Black people. Table 3.1 shows the demographic characteristics of the predicted Black and White people. Black people are relatively lower educated, less likely to be married, have lower annual family income, and less likely to live in Medicaid expansion states.

Using this sample, we estimate the ACA eligibility for Black and White people who lost ESI due to a Covid-19 layoff based on their family income, state of residence, and family

[51] The ACS does not provide information about ESI policyholders. Following (Garfield et al., 2020), we define a family-wide loss of ESI if no other person (spouse) earned more than $50,000 per year and usually worked at least 30 hours per week, who may obtain ESI through his/her own employee.

status.[52] Following recent studies, we assume all laid off workers will apply and receive regular state unemployment benefits and federal unemployment compensation. We also assume they will stay unemployed for the rest of the year.

While the $600 per week FPUC ended in July is universally applied to all unemployed people, regular state unemployment benefits differ across states. Each state has its own rule to calculate weekly unemployment benefits. For example, some states define the weekly benefits as 1/26 of average quarterly earnings in the two highest paid quarters of the base period. Some states define it as 50 percent of average weekly wage during the base period, and some states define it as 4 percent of the average quarterly earnings multiplied by 1.2075, etc. However, no matter how definitions vary, the overall weekly unemployment benefits is equivalent to about 50 percent of average weekly wage during the base period.[53] Each state also has a cap on the maximum weekly benefits and the maximum number of weeks of benefits. For example, maximum state unemployment benefits can vary from $247/week in Louisiana and $713/week in New Jersey. The majority of states provide a maximum of 26 weeks of unemployment benefits. Table C.2 shows the maximum weekly unemployment benefits and number of weeks by state. We calculate the state unemployment benefits for workers losing jobs following the above-

[52] Note, there are cases that people have intermarriage (marrying someone with different race/ethnicity). According to Livingston and Brown (2017), intermarriage has raised steadily in the past fifty years, from 3 percent to 10 percent among all married people. However, the trend varies by race/ethnicity and education. For example, Hispanic and Asian people are more likely to have intermarriage than Black and White people, and higher educated people are increasingly more likely to have intermarriage than lower educated people. In our sample, less than 2 percent of families have intermarriage.

[53] The weekly unemployment benefits can vary from 40 percent of average weekly wage in Wisconsin to 57 percent in Hawaii, based on author's calculation.

mentioned specific state rules.[54] The calculated average state unemployment benefits is about $332/week for White workers losing job and $299/week for Black workers losing job.[55]

Medicaid eligibility is based on current monthly family income. Thus, family income used to determine Medicaid eligibility is the monthly state unemployment benefits for the worker losing job plus regular monthly income from other family members. Eligibility for premium subsidy is based on projected annual family income, and the FPUC is included. Thus, family income used to determine premium subsidy eligibility is regular family income from January to March 2020, plus monthly state unemployment and FPUC benefits for the worker losing job and regular monthly income from other family members from April to December 2020. When estimating the ACA eligibility in 2021 when newly unemployed workers exhaust all unemployment compensation and are still unemployed, we replace their income with zero when calculating family income.

We calculate the ratio of family income to the FPL based on family size to determine the ACA eligibility. Nonelderly adults are defined to be Medicaid eligible if their family income (less FPUC) is below the income eligibility cut-off in the state of residence. Income eligibility cut-offs of the Medicaid program for children (or CHIP) are higher than adults. We define the Medicaid eligibility of children based on the states' CHIP income eligibility cut-offs. People who are not eligible for Medicaid with family income between 100-400 percent of the FPL are eligible for premium subsidy. Within this group, people with income below 250 percent of the FPL are additionally eligible for cost-sharing reductions.

[54] However, there may be some measurement errors. For example, I assume workers earn the same amount in each quarter by dividing their reported annual income by 4. This will underestimate the unemployment benefits in the case that states define unemployment benefits based on two highest paid quarters.

[55] National average weekly unemployment benefits were about $333 in April 2020. https://www.cbpp.org/research/economy/policy-basics-unemployment-insurance

3.4 Results

Figure 3.1 to Figure 3.3 show the estimated ACA eligibility for Black and White people we predict to have lost ESI under three scenarios: no FPUC, current level of FPUC, and when all unemployment compensation is exhausted in 2021. The first two columns summarize the percentages of Black and White people who are ACA eligible, and the second two columns summarize those who are ACA ineligible.

If the FPUC had not provided any income (results shown in Figure 3.1), White people who lost ESI are more likely than Black people to be eligible for Medicaid (50.2 percent vs. 46.8 percent) and premium subsidy (38.5 percent vs. 37.5 percent). However, White people are less likely than Black people to be eligible for additional cost-sharing reductions (24.1 percent vs. 27.3 percent). Remaining people are ACA ineligible, either because they fall in the coverage gap or because they have income beyond the premium subsidy cut-off at 400 percent of the FPL. White people would be much less likely than Black people to fall in the coverage gap (8.3 percent vs. 14.1 percent), but more likely to have income beyond the premium subsidy cut-off (3 percent vs. 1.6 percent). The overall proportion of being ACA ineligible is lower among White people than Black people (11.3 percent vs. 15.7 percent). In other words, without any government intervention, Black people who lost ESI during Covid-19 would be more likely to receive no government assistance in health insurance than White people. This is mainly caused by large disparity in the probability of falling in the coverage gap. People in coverage gap are low-income and can barely afford private health insurance, thus are very likely to be uninsured. Hence, Covid-19 could widen the existing Black-White disparity in health insurance coverage, given the existing gaps in the ACA.

So far, no federal legislation related to Covid-19 directly protects people who lost ESI from being uninsured through improving health insurance policy. However, the $600 per week FPUC ended in July may have inadvertently affected the ACA eligibility. By increasing income, the FPUC can increase access to ACA premium subsidy for people who would fall in the coverage gap without FPUC, but it also can reduce access to subsidy for other people, by making them to earn-out of the cost-sharing range, or causing them to earn-out of the entire subsidy range. Figure 3.2 shows the estimated ACA eligibility under the current FPUC. Since FPUC is not included in determining Medicaid eligibility, the percentages of Black and White people being Medicaid eligible are the same as Figure 3.1. The current FPUC does affect the premium subsidy eligibility and thus the overall ACA eligibility. The overall proportions of White and Black people being ACA ineligible become much smaller when people received FPUC. The percentage of being ACA ineligible is even higher among White people than Black people (8.2 percent vs. 5.8 percent). Hence, the current FPUC reduced the potential increase in Black-White disparity in health insurance coverage mentioned above, and may even have narrowed the existing Black-White gap in health insurance coverage. This can be explained by two factors. On the one hand, FPUC moved the majority of people who would fall in the coverage gap to become premium subsidy eligible. With current FPUC, the percentage falling in the coverage gap dropped from 8.3 percent to 0.9 percent among White people and from 14.1 percent to 1.3 percent among Black people, closing the large Black-White disparity in the likelihood of being in the coverage gap. On the other hand, FPUC made more White people earn-out of the entire subsidy range by increasing their income beyond 400 percent of the FPL, as the percentage of having income above the cut-off increased by 4.3 percent (from 3 percent to 7.3 percent) among

White people who lost ESI and by 2.9 percent (from 1.6 percent to 4.5 percent) among Black people.

Overall, Figure 3.2 show that, the percentages of being ACA ineligible among Black and White people who lost ESI are smaller when they received FPUC. The major reason for being ineligible is having income above the premium subsidy cut-off at 400 percent of the FPL, and White people who lost ESI are more likely to face this situation. This is different from the previous scenario, where the major reason for being ACA ineligible is falling into the coverage gap. People with income above 400 percent of the FPL are comparatively more likely to purchase unsubsidized private health insurance than people in the coverage gap – those with low income. Thus, it is less conclusive whether the current FPUC, which causes larger proportion of White people who lost ESI be ACA ineligible, has narrowed the existing Black-White disparity in health insurance coverage. [56] Nevertheless, it is clear that the current FPUC can reduce the potential increase in Black-White disparity in health insurance coverage during Covid-19 if there was no FPUC.

As mentioned earlier, FPUC can also make people to earn-out of the cost-sharing range by increasing their income above the 250 percent of the FPL. We do not find the current FPUC impacts Black and White people differently in this case. About 15 percent of White people and 14.6 percent of Black people lost cost-sharing when receiving FPUC.[57]

[56] Unsubsidized private health insurance can still be very expensive for middle-income families. If all people with income above 400 percent of the FPL choose to be uninsured, with more White people than Black people who lost ESI facing this situation under current FPUC, the existing Black-White disparity in health insurance coverage may be narrowed. However, we have to interpret this result with caution, as reducing eligibility for White people is not the intended way to reduce racial/ethnic disparity.

[57] The percentage of White people lost cost-sharing is calculated as 0.241 + (0.083 – 0.009) - 0.165 = 0.15. The percentage of Black people lost cost-sharing is calculated as 0.273 + (0.141 – 0.013) - 0.255 = 0.146.

The current FPUC program can temporarily help address the issue of the potential increasing Black-White disparity in health insurance coverage during Covid-19. However, given the uncertainties of future economy recovery and continuity of government unemployment compensation, it is important to understand what would happen in the worst case that unemployed people exhaust all unemployment compensation and cannot return to work next year. In Figure 3.3, we estimate the ACA eligibility in 2021, assuming people who lost job and ESI during Covid-19 remain unemployed and receive no more unemployment compensation. With further lower income because of not receiving state and federal unemployment benefits anymore, many more people will be eligible for Medicaid. About 71.8 percent of White people who lost ESI will be Medicaid eligible, while 67.2 percent of Black people will be Medicaid eligible. White people will also be more likely than Black people to be eligible for premium subsidy (12.9 percent vs. 8.7 percent). The majority of people eligible for premium subsidy are eligible for additional cost-sharing reductions. A substantial amount of people will be ACA ineligible because of falling in the coverage gap. The percentage is much lower among White people than Black people (15.1 percent vs. 24 percent). Nearly zero percent of people will be ACA ineligible because of having income above 400 percent of the FPL. To sum up, if the existing gaps in the ACA are not addressed and people exhaust all unemployment compensation in the worst case, the Black-White disparity in health insurance coverage will be significantly exacerbated.

The above estimations are based on a sample of people whom we assume to lose ESI because a family member is in the top 10 percentile of our predicted job loss distribution. We additionally check the sensitivity of our results with different assumptions. Figure 3.4 shows the calculated Black-White gap in ACA ineligibility if we limit the sample to people we assume to

lose ESI because a family member is in the top 10, 15, and 20 percentiles of our predicted job loss distribution. As can be seen, all three scenarios show similar results: Black people who lost ESI are more likely than White people to be ACA ineligible when there was no FPUC, less likely than White people to be ACA ineligible with the current FPUC, and much more likely than White people to be ACA ineligible in 2021 if they exhaust all unemployment compensation.

3.5 Conclusion/Discussion

A massive number of people have lost ESI due to job loss during the Covid-19 recession. Having access to health insurance coverage is essential to protect against barriers of seeking healthcare, such as testing and treatment. However, current federal legislations to remediate the economic consequences of Covid-19 have not directly addressed the issue of improving health insurance coverage among the newly unemployed and their families.

The ACA extends affordable coverage options, through Medicaid and premium subsidy. However, there have been relevant concerns that current provisions of the ACA leave some people outside the reach of ACA benefits. Because of eligibility rules, federal unemployment compensation (FPUC) could inadvertently alter access to ACA benefits. In this study, we investigate how Covid-19, existing gaps in the ACA, and FPUC interact with each other to affect Black-White disparity in the ACA eligibility or access to health insurance coverage. Given the racial/ethnic disparities in health insurance coverage and healthcare access in the U.S. over past decades, and the disproportionate impact of Covid-19 on the health of Black people, it is important to understand the Black-White differences in obtaining alternative health insurance after losing ESI.

Our findings show that, without FPUC, Black people who lost ESI due to a Covid-19 layoff would be more likely than White people to receive no ACA benefits, enlarging the existing Black-White disparity in health insurance coverage. This is mainly caused by the gaps in current Medicaid eligibility, as more Black people live in states with less generous Medicaid program and thus fall in the coverage gap. Current FPUC that ended in July reduced the potential increase in Black-White disparity in health insurance coverage, by increasing access to premium subsidy for people who would fall in the coverage gap.

We further explore what would happen to Black-White disparity in health insurance coverage in early 2021 if people lost jobs and ESIs during Covid-19 remain unemployed and exhaust all unemployment compensation. We find that Black people are much more likely than White people to be ACA ineligible because of falling in the coverage gap, which can substantially widen the existing Black-White disparity in health insurance coverage. Given the uncertainties in future economy recovery and government decisions about unemployment compensation, closing the gaps in the current ACA is very important for building racial/ethnic equity in access to health insurance and healthcare.

Recent studies have proposed several solutions to bolster ACA health insurance coverage to protect more people from economic and health hardship (e.g., Blumberg et al., 2020a; Blumberg et al., 2020b; Straw et al, 2020). Expanding Medicaid in the non-expansion states has been strongly recommended. After the ACA went into effect, Black-White disparity in health insurance coverage has been narrowed but still persists. The major reason is that the unequal expansion of Medicaid across states leaves more Black people in the coverage gap (Buchmueller et al., 2016). In our study, we show that Covid-19 could widen the Black-White disparity in health insurance coverage when no FPUC is extended to people who lost jobs, and the disparity

may increase more in 2021 if people cannot return to work and unemployment compensation is exhausted. Recent studies find Black people are less likely to return to work after being laid off during Covid-19 (e.g., Montenovo et al. 2020). This evidence indicates that expanding Medicaid in the non-expansion states is very important for allowing Black people to have more equal access to health insurance as White people and avoiding Black-White disparity in health insurance coverage to increase, particularly under the pandemic. However, expanding Medicaid is the states' decision. Among states that have not expanded Medicaid by 2020, only the state of Nebraska decided to expand Medicaid recently. Another way to help more Black people at-risk of falling in the coverage gap is to expand the income range for premium subsidy eligibility, which is the federal government's decision. For states that refuse to expand Medicaid, extending premium subsidy to people in the coverage gap can also allow more Black people to have access to affordable health insurance, and reduce the Black-White disparity in health insurance coverage.

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Table 3.1. Summary of demographic characteristics of Black and White people who lost ESI

	White people		Black people	
	Mean	Sd	Mean	Sd
Male	0.47	0.50	0.47	0.50
Age	30.36	15.54	31.33	15.58
Education				
Less than high school	0.15	0.36	0.20	0.40
High school/GED	0.31	0.46	0.33	0.47
Some college	0.42	0.49	0.41	0.49
College and above	0.12	0.32	0.07	0.25
Married	0.24	0.43	0.17	0.38
Family size	1.79	1.24	1.82	1.24
Annual family income in the past year	$40,927	$41,476	$35,592	$33,311
Live in Medicaid expansion states	0.76	0.43	0.61	0.49
Obs	54351		13374	

Note: Data is from ACS 2018. Sample includes predicted nonelderly Black and White people who lost ESI due to a Covid-19 layoff. Results are weighted by ACS person weight.

Figure 3.1. Estimated ACA eligibility with no FPUC

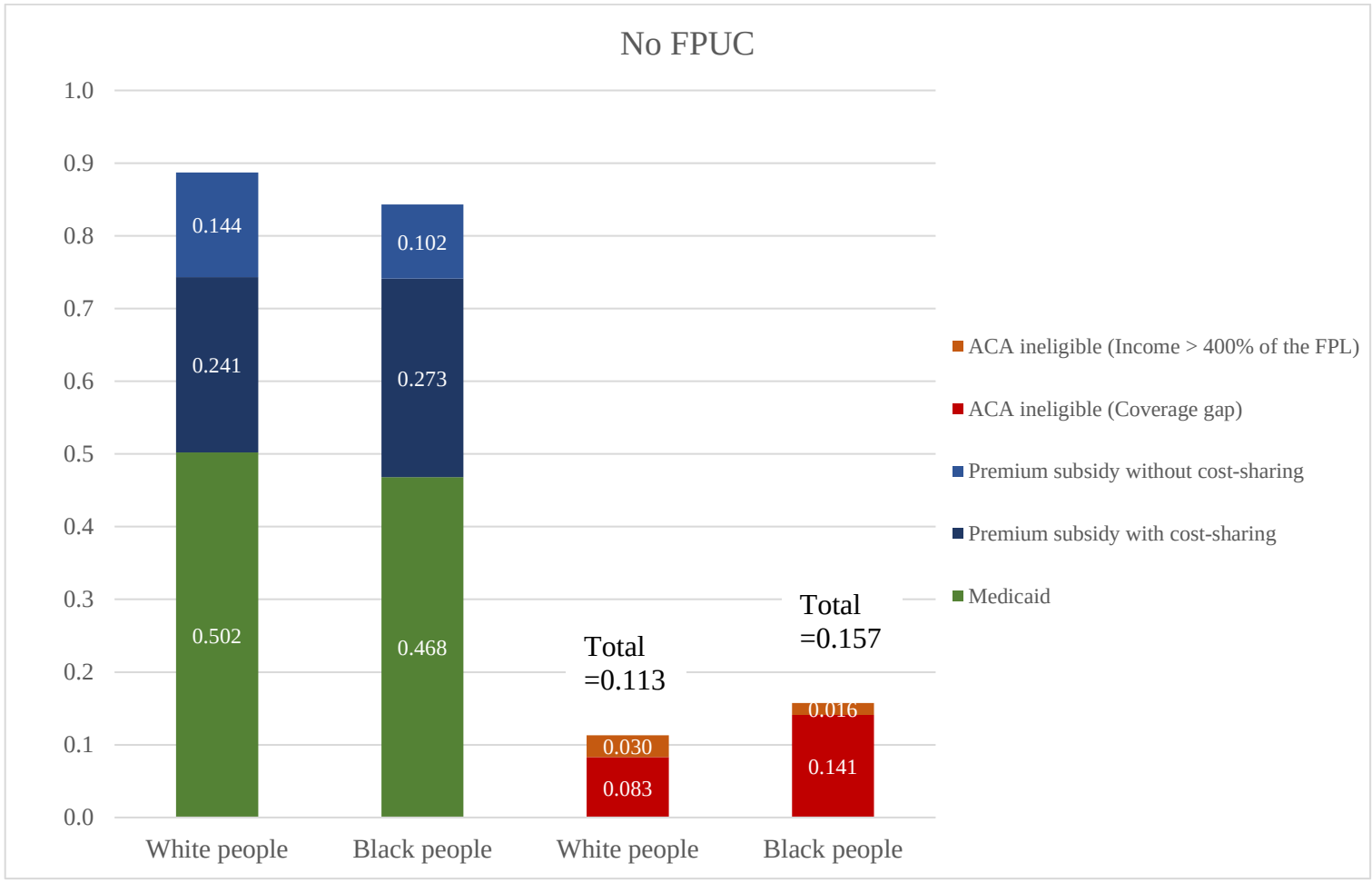


Figure 3.2. Estimated ACA eligibility under the current FPUC

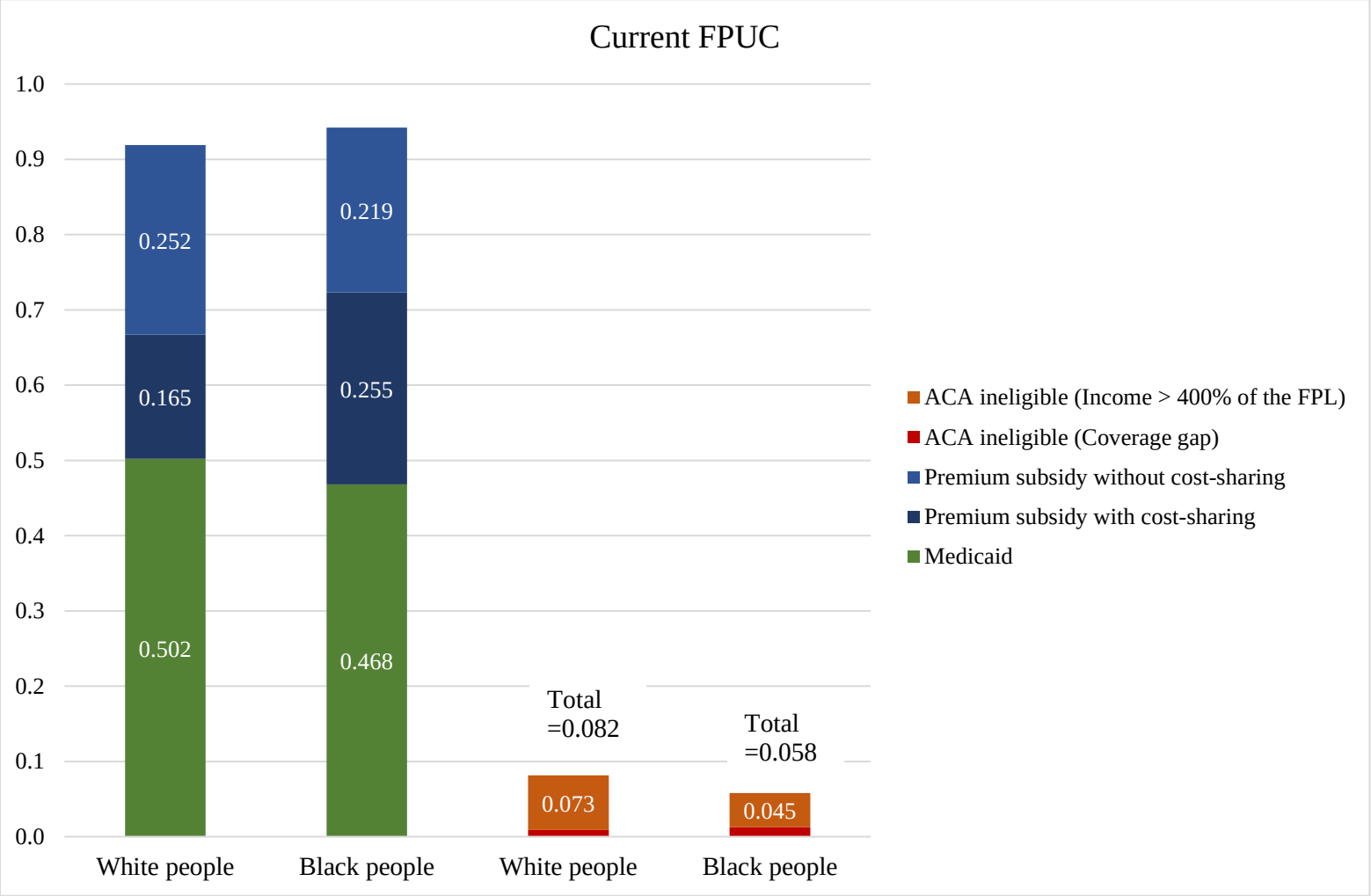


Figure 3.3. Estimated ACA eligibility in year 2021 if unemployment compensation is exhausted

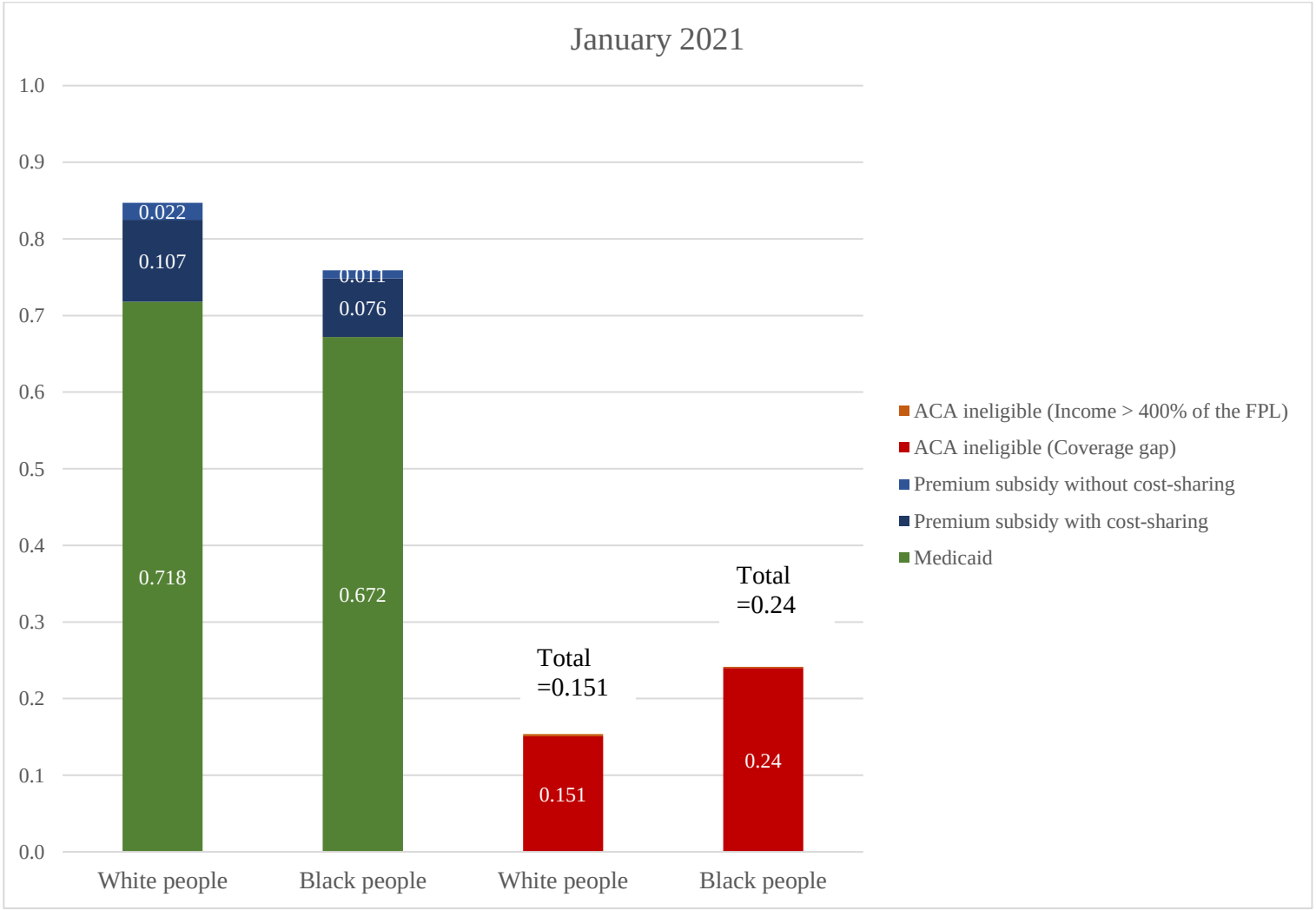
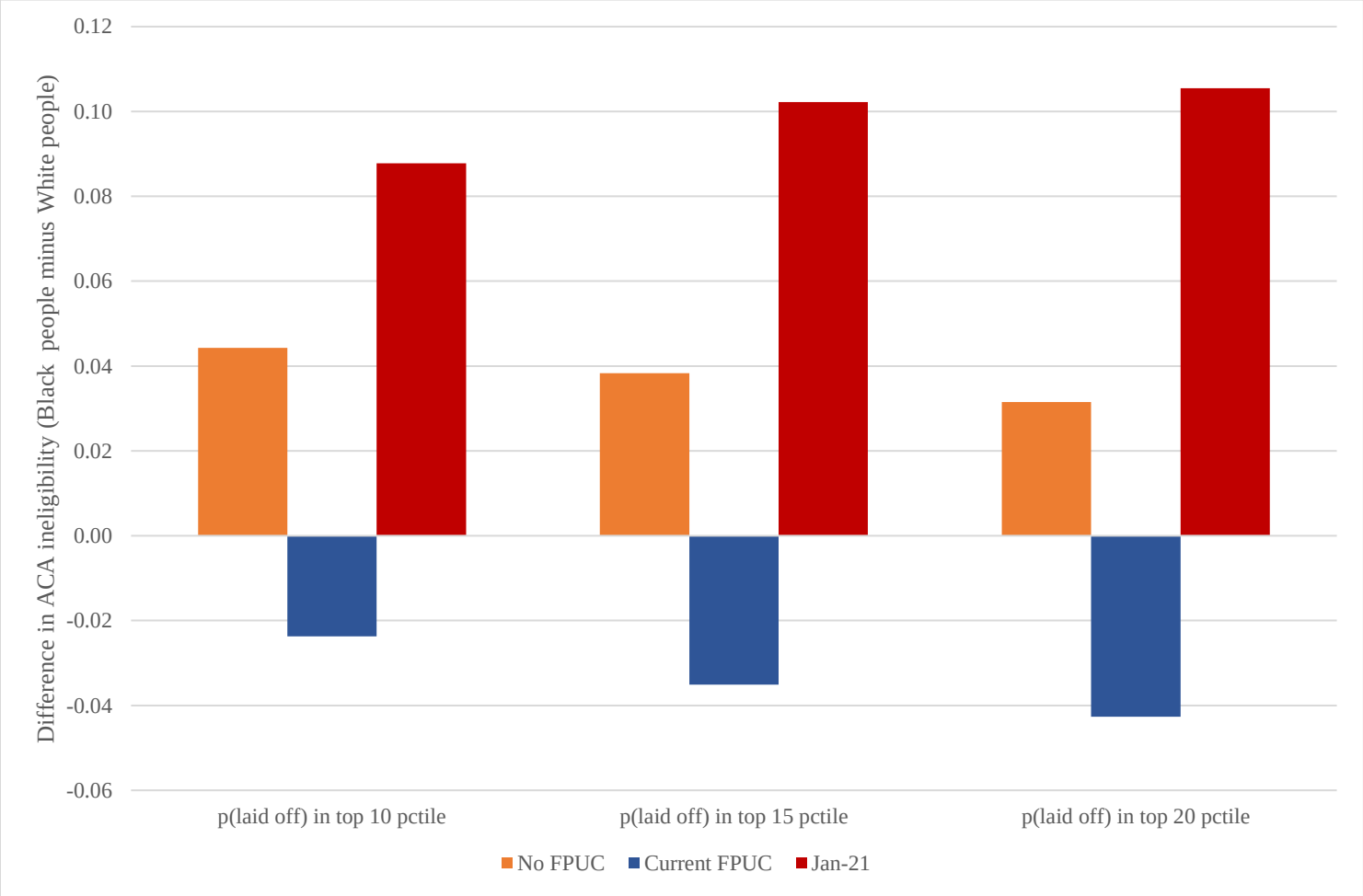


Figure 3.4. Estimated Black-White disparity in ACA eligibility based on different assumptions



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Appendix A: Chapter 1 Appendix

Table A.1. Medicaid income eligibility cutoffs and states' adoption status

	2013 thresholds		Adopt by Jan. 2014?
	Parents	Childless adults	
Alabama	0.23	0	N
Alaska	0.78	0	9/1/15
Arizona	1.06	1	Y
Arkansas	0.16	0	Y
California	1.06	0	Y
Colorado	1.06	0.2	Y
Connecticut	0.91	0.7	Y
Delaware	1.2	1.1	Y
District of Columbia	2.06	2.11	Y
Florida	0.56	0	N
Georgia	0.48	0	N
Hawaii	1.38	1	Y
Idaho	0.37	0	N
Illinois	1.39	0	Y
Indiana	0.24	0	2/1/15
Iowa	0.8	0	Y
Kansas	0.31	0	N
Kentucky	0.6	0	Y
Louisiana	0.25	0	7/1/16
Maine	2	0	N
Maryland	1.2	0	Y
Massachusetts	1.33	0	Y
Michigan	0.64	0	4/1/14
Minnesota	2.15	0.75	Y
Mississippi	0.29	0	N
Missouri	0.35	0	N
Montana	0.55	0	1/1/16
Nebraska	0.58	0	N
Nevada	0.84	0	Y
New Hampshire	0.49	0	8/15/14
New Jersey	2	0	Y
New Mexico	0.85	0	Y
New York	1.5	1	Y
North Carolina	0.59	0	N
North Dakota	0.59	0	Y
Ohio	0.9	0	Y
Oklahoma	0.5	0	N
Oregon	0.4	0	Y
Pennsylvania	0.58	0	1/1/15
Rhode Island	1.81	0	Y
South Carolina	0.9	0	N
South Dakota	0.5	0	N
Tennessee	1.22	0	N
Texas	0.26	0	N
Utah	0.44	0	N
Vermont	1.91	1.6	Y
Virginia	0.3	0	N
Washington	0.71	0	Y
West Virginia	0.31	0	Y
Wisconsin	2	0	N
Wyoming	0.5	0	N

Appendix B: Chapter 2 Appendix

Table B.1. Contribution schedule for the ACA premium tax credits

FPL	Contribution (as % of income)	
	Medicaid non-expansion states	Medicaid expansion states
<100%	Medicaid No premium subsidy (“coverage gap”)	Medicaid
100%-138%	2%	Medicaid
138%-150%	3-4%	3-4%
150%-200%	4-6.3%	4-6.3%
200%-250%	6.3-8.05%	6.3-8.05%
250%-300%	8.05-9.5%	8.05-9.5%
300-400%	9.5%	9.5%

Table B.2. Marriage penalties for two childless single adults with hypothetical income levels

			Married				Unmarried				
		Income	Income to the FPL	Medi caid	Premium tax credits	Total premium payments	Income to the FPL	Medi caid	Premium tax credits	Total premium payments	**Marriage penalties**
Panel A: In Medicaid expansion states											
Case 1	A	$10,000	127%	yes	no	$0	86%	yes	no	$0	$0
	B	$10,000					86%	yes	no	$0	
	Total	$20,000								$0	
Case 2	A	$10,000	254%	no	yes	$3,400	86%	yes	no	$0	$850
	B	$30,000					257%	no	yes	$2,550	
	Total	$40,000								$2,550	
Case 3	A	$10,000	381%	no	yes	$5,700	86%	yes	no	$0	-$300
	B	$50,000					428%	no	no	$6,000	
	Total	$60,000								$6,000	
Case 4	A	$10,000	445%	no	no	$12,000	86%	yes	no	$0	$6,000
	B	$60,000					514%	no	no	$6,000	
	Total	$70,000								$6,000	
Case 5	A	$20,000	254%	no	yes	$3,400	171%	no	yes	$1,000	$1,400
	B	$20,000					171%	no	yes	$1,000	
	Total	$40,000								$2,000	
Case 6	A	$20,000	381%	no	yes	$5,700	171%	no	yes	$1,000	$900
	B	$40,000					343%	no	yes	$3,800	
	Total	$60,000								$4,800	
Panel B: In Medicaid non-expansion states											
Case 7	A	$10,000	127%	no	yes	$400	86%	no	no	$6,000	-$11,600
	B	$10,000					86%	no	no	$6,000	
	Total	$20,000								$12,000	
Case 8	A	$10,000	254%	no	yes	$3,400	86%	no	no	$0	$850
	B	$30,000					257%	no	yes	$2,550	
	Total	$40,000								$2,550	
Case 9	A	$10,000	381%	no	yes	$5,700	86%	no	no	$6,000	-$6,300
	B	$50,000					428%	no	no	$6,000	
	Total	$60,000								$12,000	
Case 10	A	$10,000	445%	no	no	$12,000	86%	no	no	$6,000	$0
	B	$60,000					514%	no	no	$6,000	
	Total	$70,000								$12,000	
Case 11	A	$20,000	254%	no	yes	$3,400	171%	no	yes	$1,000	$1,400
	B	$20,000					171%	no	yes	$1,000	
	Total	$40,000								$2,000	

Table B.3. Marriage penalties for childless individual A marrying individual B with different number of children

		Married				**Unmarried**				**Marriage penalties**
	Income	Income to the FPL	Medi caid	Premium tax credits	Total premium payments	Income to the FPL	Medi caid	Premium tax credits	Total premium payments	
Panel A:										
A	$30,000					257%	no	yes	$2,415	
B (with 1child)	$30,000	381%	no	yes	$5,700	257%	no	yes	$2,415	$870
Total	$60,000								$4,830	
A	$30,000					257%	no	yes	$2,415	
B (with 1 child)	$30,000	303%	no	yes	$5,700	191%	no	yes	$1,800	$1,485
Total	$60,000								$4,215	
A	$30,000					257%	no	yes	$2,415	
B (with 2 children)	$30,000	252%	no	yes	$4,860	152%	no	yes	$1,200	$1,245
Total	$60,000								$3,615	
Panel B:										
A	$50,000					428%	no	no	$6,000	
B (with 0 child)	$30,000	509%	no	no	$12,000	257%	no	yes	$2,430	$3,570
Total	$80,000								$8,430	
A	$50,000					428%	no	no	$6,000	
B (with 1 children)	$30,000	404%	no	no	$14,000	191%	no	yes	$1,890	$6,110
Total	$80,000								$7,890	
A	$50,000					428%	no	no	$6,000	
B (with 2 children)	$30,000	335%	no	yes	$7,600	152%	no	yes	$1,200	$400
Total	$80,000								$7,200	

Note: Premium of the individual private health insurance is assumed to be $6,000 for adults and $2,000 for children.

Table B.4. Characteristics of predicted partner of single individuals

	Single male	Single female
Predicted partner's income	$30,205	$35,396
Predicted partner's age	45.47	46.34
Number of predicted partner's children	0.49	0.18
Obs	855849	1144556

Note: Data from the ACS 2011-2017.

Figure B.1. Marriage penalties for all possible combinations of income of each partner, assuming both partners are 30-year-old and living in a Medicaid expansion state

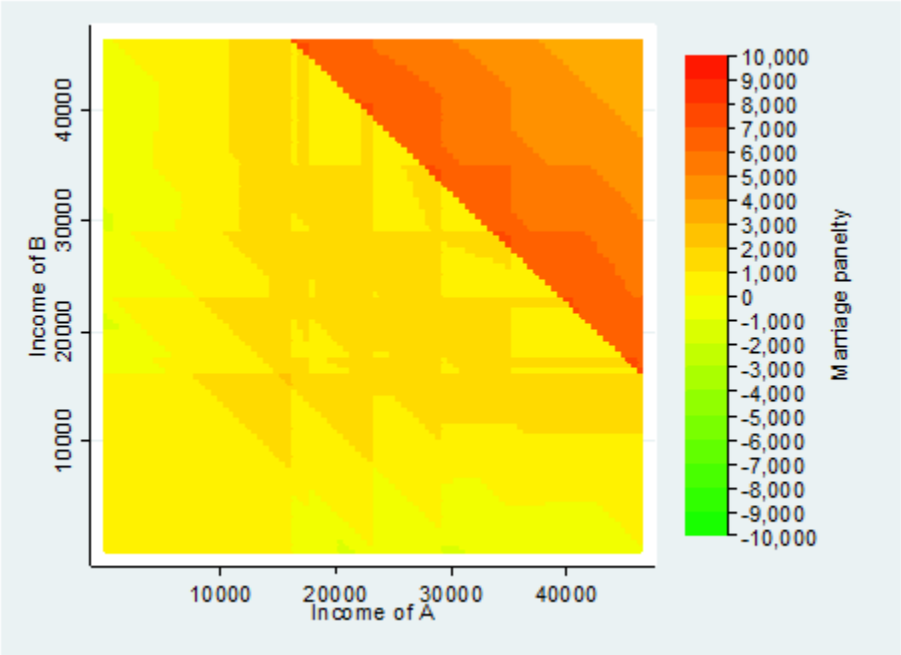


Figure B.2. Marriage penalties for all possible combinations of income of each partner, assuming both partners are 30-year-old and living in a Medicaid non-expansion state

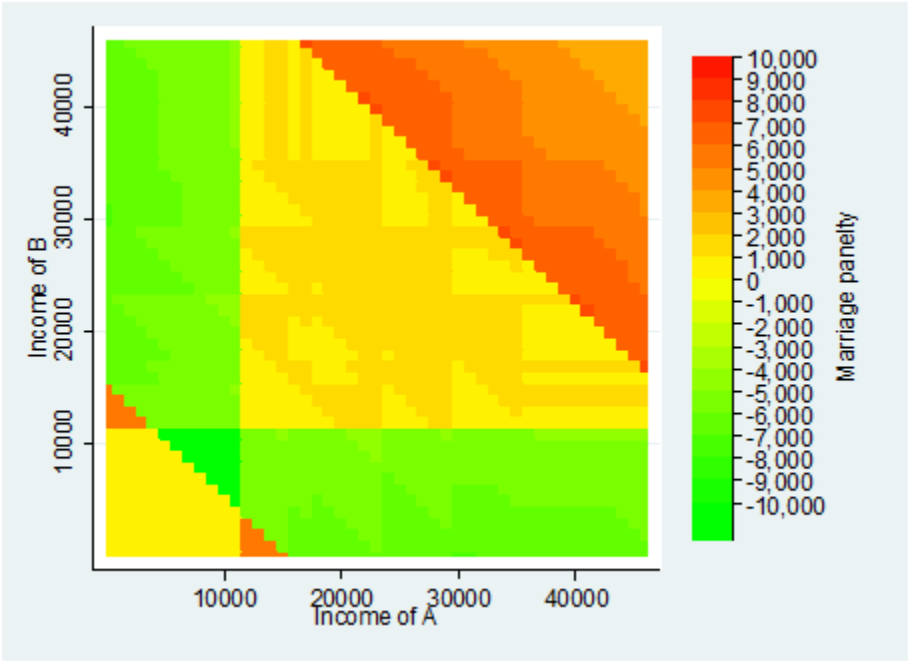
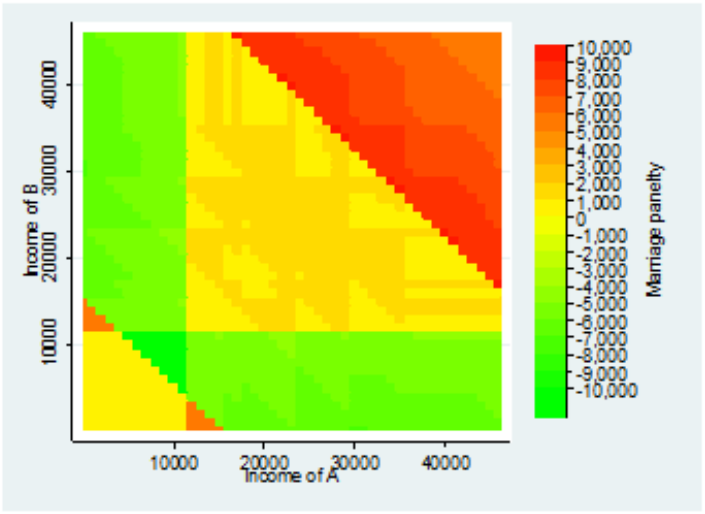
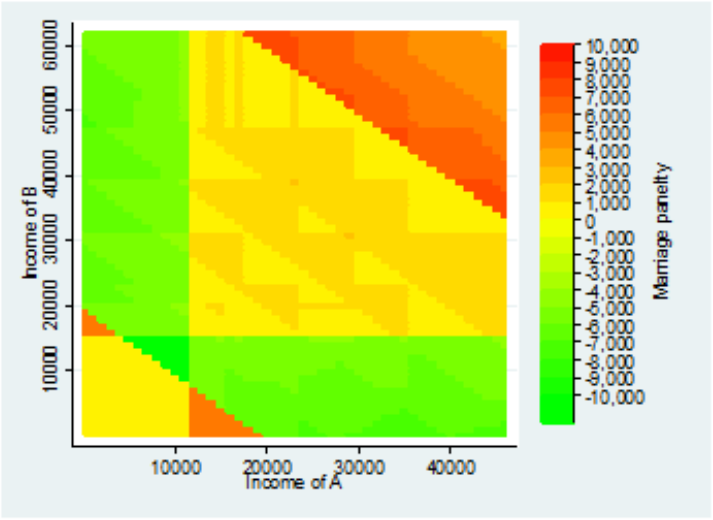


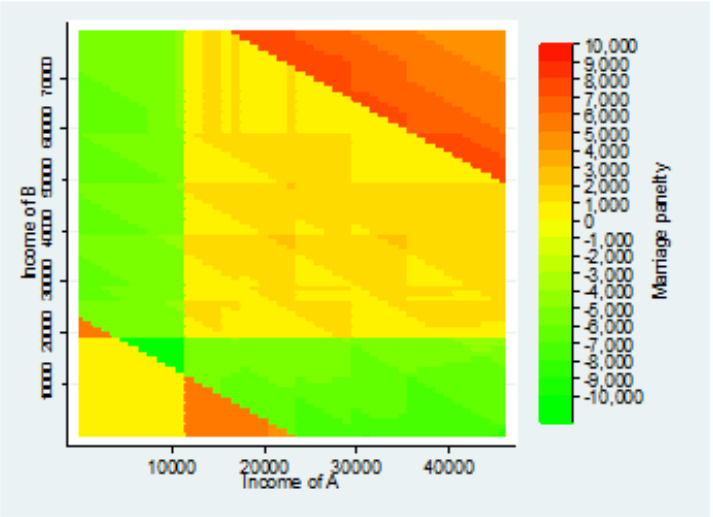
Figure B.3. Marriage penalties by the number of partner's children



(B has 0 kid)

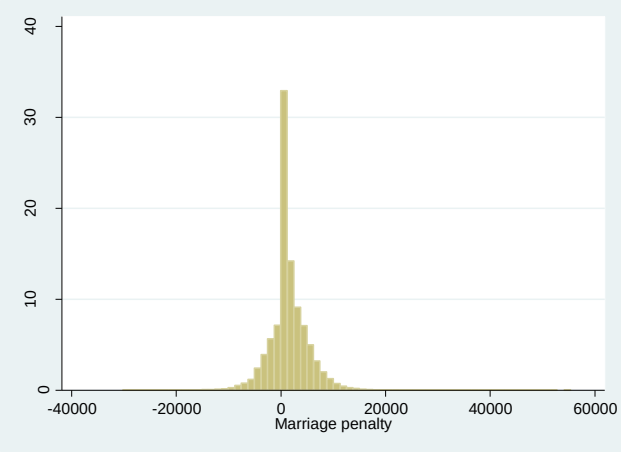


(B has 1 kid)

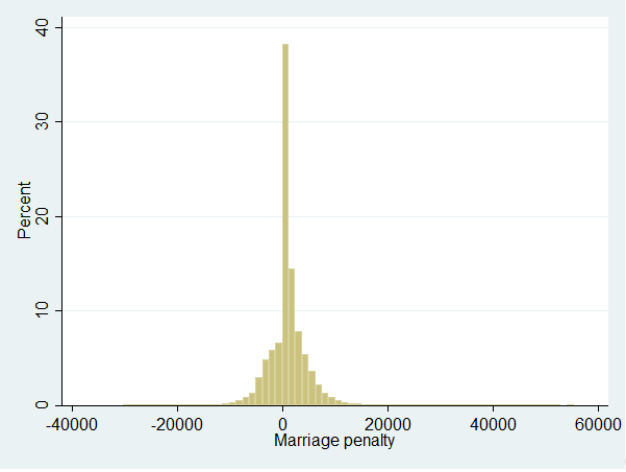


(B has 2 kids)

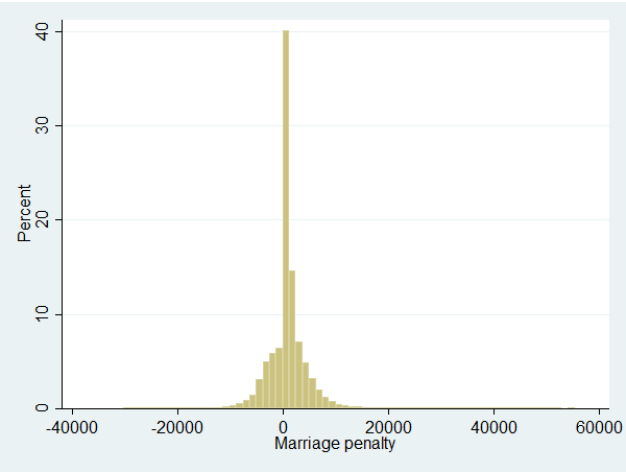
Figure B.4. The distributions of estimated ACA marriage penalties using different assumptions about the assignment of children if married couples divorce.



(assign children to women; mean = $1,335)



(assign children to men; mean = $848)



(assign children to higher earner; mean = $737)

Appendix C: Chapter 3 Appendix

Table C.1. Relationship between individual characteristics and recent unemployment

Variable	losing job between March and May 2020 (log odds)
Age	-0.04468***
	(0.008)
Age square	0.00046***
	(0.000)
Male	-0.24932***
	(0.041)
Race (reference = White)	
Black	0.24995***
	(0.070)
Hispanic	0.22291**
	(0.099)
Asian/other	0.41444***
	(0.108)
Education (reference: less than high school)	
High school/GED	0.09155
	(0.071)
Some college	0.10223
	(0.068)
College and above	-0.27370***
	(0.091)
Married	-0.19485***
	(0.042)
Worked in non-essential business	0.73252***
	(0.088)
Obs	45795

Notes: Data is from 2020 CPS May. Model also controls for 25 dummies of occupation group. 18 dummies of industry group, and state fixed effects. Standard errors in parentheses are clustered at state level. * p<0.05 ** p<0.01 *** p<0.001.

Table C.2. Regular maximum weekly unemployment benefits and maximum weeks by state

State	Max.weekly benefit	Max. weeks	State	Max.weekly benefit	Max. weeks
Alabama	$275	14	Nebraska	$440	26
Alaska	$370	26	Nevada	$469	26
Arizona	$240	26	New Hampshire	$427	26
Arkansas	$451	16	New Jersey	$713	26
California	$450	26	New Mexico	$511	26
Colorado	$618	26	New York	$504	26
Connecticut	$649	26	North Carolina	$350	12
Delaware	$400	26	North Dakota	$618	26
District of Columbia	$425	26	Ohio	$480	26
Florida	$275	12	Oklahoma	$539	26
Georgia	$365	26	Oregon	$648	26
Hawaii	$648	26	Pennsylvania	$573	26
Idaho	$405	26	Rhode Island	$586	26
Illinois	$484	26	South Carolina	$326	26
Indiana	$390	26	South Dakota	$414	20
Iowa	$511	26	Tennessee	$275	26
Kansas	$488	26	Texas	$521	26
Kentucky	$552	26	Utah	$580	26
Louisiana	$247	26	Vermont	$513	26
Maine	$445	26	Virginia	$378	26
Maryland	$430	26	Washington	$844	26
Massachusetts	$823	26	West Virginia	$424	26
Michigan	$362	20	Wisconsin	$370	26
Minnesota	$740	26	Wyoming	$508	26
Mississippi	$235	26			
Missouri	$320	20			
Montana	$552	28			

Source: Collected from each state's unemployment benefit website. Under CARES Act in response to Covid-19, all states extend extra 13 weeks of unemployment benefits through federally funded Pandemic Emergency Unemployment Assistance (PEUC) benefits.