The Relationship Between Health Insurance and Early Retirement: Evidence from the Affordable Care Act

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Abstract

This paper investigates the effect of the Affordable Care Act's (ACA) Medicaid expansion on the retirement decision of low-educated adults aged 55-64. Using data from the American Community Survey, I employ a difference-in-differences strategy that exploits the timing and expansion decisions of states for adults without dependent children ("childless adults"). I find that the expansions increase Medicaid enrollment for both men and women. The estimates also suggest that the expansions and Medicaid enrollment result in women retiring early, whereas there is no significant change in the retirement behavior of men. These findings imply that the effect of health insurance on women's retirement decisions may depend on men's labor market responses to health insurance.

Keywords: Medicaid, retirement, Affordable Care Act, job lock
JEL: I18, J18, J26

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1. Introduction

The Medicaid program is a means-tested health insurance program that expanded in recent decades to increase access to health coverage among low-income families, mainly those with children. The Affordable Care Act's (ACA) Medicaid expansion specifically targeted low-income adults without dependent children ("childless adults") – a group that had limited access to Medicaid prior to the ACA. Although the Supreme Court decision made the ACA's Medicaid expansion optional for states, the majority of the states expanded coverage to individuals below 138% of the federal poverty level (FPL), which is about $22,715 for a family of two in 2018.[1] Starting with the early expansions in 2010, the uninsured rate reached a record low in 2015 (Sommers, Kenney, and Epstein, 2014, Cohen, Martinez, and Zammitti, 2016). Given the availability of Medicaid as an alternative to employer-sponsored health insurance (ESHI), understanding the relationship between the ACA's Medicaid expansion and labor supply has become crucial in the existing literature.[2]

A question that follows the premise that the ACA's Medicaid expansion disincentivizes work is whether the availability of Medicaid affects the retirement decision of workers. There has been a steady decline in retiree benefits offered by large firms, defined as 200 or more workers, from 40% in 1999 to 25% in 2017.[3] If labor market behavior is affected by the availability of health insurance, then early retirement, defined as leaving the labor force before the age of 65, could be used to test for the job-lock effect under the ACA's Medicaid expansion.[4] Although there is a developed body of literature showing the effect of retiree health insurance (RHI) on retirement (see Section 2 for a detailed literature review), less is known about the relationship between Medicaid and retirement. This study is a contribution to the existing retirement literature by exploring the effect of the ACA's Medicaid expansion on early retirement of childless adults aged 55-64.

In this paper, I introduce a simple static model to show the effect of health insurance on leisure. I consider cases where Medicaid enrollment is exogenously determined, and

[1] As of January 2018, there are 33 expansion states and 18 non-expansion states.

[2] There are a number of studies that investigate the relationship between labor market outcomes and the ACA's Medicaid expansion (Kaestner et al., 2017, Duggan, Goda, and Jackson, 2017, Aslim, 2016, Gooptu et al., 2016, Kim, 2016, Leung and Mas, 2016).

[3] See Figure 11.1, 2017 Employer Health Benefits Annual Survey, Kaiser Family Foundation and Health Research & Educational Trust, retrieved May 1, 2017 from: `https://www.kff.org/health-costs/report/2017-employer-health-benefits-survey/`.

[4] For example, Bailey and Chorniy (2016) investigate the job lock effect using the dependent coverage mandate as a natural experiment.

when individuals self-select into Medicaid by reducing working hours. While health insurance generates a positive income effect by reducing medical expenses, it also increases the time spent on leisure and work through fewer sick days. Assuming leisure is a normal good, both cases imply that Medicaid enrollment and leisure have a positive relationship. I test this prediction using data from the 2010-2016 Public Use Micro Samples (PUMS) of the American Community Survey (ACS). Additionally, a difference-in-differences (DD) approach is employed to exploit the timing and expansion decision of states under the ACA. I estimate the changes in Medicaid enrollment and retirement for low-educated women and men, and obtain Wald estimates (or the treatment-on-treated effects) using these intention-to-treat (ITT) effects.

My findings suggest that the expansions increase Medicaid enrollment by 5 percentage points for men and 6 percentage points for women.[5] I find that the expansions do not affect the retirement decision of men, whereas women increase their probability of retirement by 0.6 percentage points.[6] Findings also show that the retirement effect is stronger for women aged 59-64. This finding implies that men and women have different responses to acquiring health insurance. In support of this, Boyle and Lahey (2016) find that financial incentives dominate leisure complementarity among spouses vis-á-vis health insurance, an effect that is found to be stronger for low-educated wives. For women, Wald estimates also yield positive effects of Medicaid enrollment on retirement, albeit relatively large. When the sample is restricted to high-educated men and women, I do not find any significant effects on retirement. These findings are robust to a number of alternative specifications and sample periods.

The paper is organized as follows: Section 2 provides a detailed summary of the literature, Section 3 introduces a theoretical framework on leisure and health insurance, Section 4 provides a background information on expansion and non-expansion states, Section 5 introduces the data, Section 6 describes the empirical methodology, Section 7 presents the results, and Section 8 concludes the paper.

[5] These estimates are similar to those found in the literature (see, for example, Kaestner et al., 2017).

[6] Using the 2006 Massachusetts health reform as a natural experiment, Heim and Lin (2017) find an increase in women's early retirement from full-time employment by 1.1 percentage points and no effect on men's retirement behavior.

2. Prior Research on Health Insurance and Early Retirement

This section provides a review of studies that explore the effects of health insurance on early retirement. Previous studies mainly focus on the availability of retiree health insurance (RHI) and its impact on early retirement. In order to make these studies comparable, I categorize them with respect to their choice of data set. I start the analysis with the papers that use the Survey of Income and Program Participation (SIPP) and the Current Population Survey (CPS). Next, I compare the studies that use the Health and Retirement Survey (HRS), which is in fact the most commonly used data set in the literature. I leave the discussion of the studies that use administrative or confidential data and those that specifically focus on the ACA for the end.

Madrian (1994a) provides earlier evidence on this issue using the 1987 National Medical Expenditure Survey (NMES) along with two modules of the Survey of Income and Program Participation (SIPP) for the panels between 1984 and 1986.[7] Using age at retirement as the outcome variable, the paper finds that individuals with RHI retire 5 to 16 months earlier than those without health insurance benefits.[8] The paper further shows sizable reductions in labor force participation vis-à-vis a reform towards universal coverage. Karoly and Rogowski (1994) also use the SIPP to investigate the effect of "continuation coverage," which is the continuation of ESHI after retirement, on the decision to retire. Karoly and Rogowski (1994) differ from Madrian (1994a) by extending the analysis to the 1988 panel of SIPP. The findings, however, are consistent in terms of showing a positive effect on the probability of retirement.

There are a number of studies that use the Health and Retirement Survey (HRS) as the main data source in a dynamic retirement framework. Marton, Woodbury, and Wolfe (2007) use three waves from the HRS between 1992-1996 and find a 55 percent increase in the probability of retirement for workers that have access to retiree health benefits.[9] They also find heterogeneous effects for men with a full-time employed wife and unmarried men. Using the same data period from the HRS, Rogowski and Karoly (2000) find an increase in the probability of retirement by 68 percent for workers with

[7]These two modules of SIPP are on education, work history, and job characteristics.

[8]The paper addresses censoring problems related to the outcome variable, as well as selection into jobs.

[9]The authors are conservative with respect to the causal interpretation of their estimates.

RHI.[10] Both studies define retirement as a transition from full-time employment to being retired in the following wave(s). The difference in the estimates, however, is argued to be driven by the differential coding of the existing control variables and/or the sample size of the last wave.

In a study that exploits a discrete-time hazard model, Marton and Woodbury (2013) estimate the effect of delayed payment contracts (in the form of retiree health benefits) on the retirement decision of workers by different age groups. When given retiree health benefits, the study shows that workers at the ages of 50 and 51 are less likely to retire than those at the ages of 60 and 61. Robinson and Clark (2010), on the other hand, use a Cox proportional hazard model to analyze the impact of RHI on the decision to separate from employment. Using eight waves from the HRS between 1992-2006, the findings suggest an increase in the likelihood of job seperation by 21.2 percent for individuals with access to RHI. Kapur and Rogowski (2011) also use the same eight waves from the HRS to study the retirement decision of women with respect to the availability of RHI.[11] The availability of RHI increases the probability of retirement by 3 percentage points and 4.8 percentage points for women in dual-earner couples and single women, respectively.

Strumpf (2010) uses the HRS not only to show the impact of RHI on early retirement, but also to analyze the changes in health care utilization, medical costs, and health outcomes. Using a probit model, she finds an increase in the probability of early retirement by 37 percent for workers with RHI. Shoven and Slavov (2014), on the other hand, exploit a multinomial logit to model the retirement decision as a transition from full-time employment to part-time employment ("partial retirement") or leaving the labor force. The sample is restricted to public sector workers who have access to group health coverage before qualifying for Medicare. The paper shows an increase in the probability of leaving full-time employment by 38 percent and 26 percent for workers in age groups 55-59 and 60-64, respectively. The findings also suggest that workers in the former group are more likely to transition into part-time employment, whereas those in the latter group are more likely to leave the labor force. Blau and Gilleskie (2001) also incorporate employment transitions using a multinomial logit model. Different from prior studies, the model accounts for a possible correlation between the unobservable factors affecting

[10] Marton, Woodbury, and Wolfe (2007) use the final release of the HRS data for wave three, whereas Rogowski and Karoly (2000) use the alpha release of the data for wave three, which has less observations.

[11] Retirement is defined as a transition from full-time employment in the baseline year to retirement at the next survey date.

health insurance and employment decisions. The results indicate an increase in the exit rate from employment by 2 percentage points when individuals share the cost of RHI with the firm, and the magnitude increases as the firm pays all of the insurance costs.

Although reduced form models are common in the existing literature, some studies use the HRS to estimate a dynamic stochastic model of retirement. Blau and Gilleskie (2006) simulate the retirement decision for multiple scenarios including the availability of RHI for those without any coverage, a system with universal health coverage, and increasing the age threshold for Medicare. The authors show differential effects for men and women, where non-employment increases for the former (3.1 percentage points) and decreases for the latter (1.8 percentage points) under RHI. Blau and Gilleskie (2008), on the other hand, show an increase in non-employment by 3.6 percentage points for men who have ESHI with retiree health benefits compared to those without retiree health benefits. It is crucial to note that Blau and Gilleskie (2006) and Blau and Gilleskie (2008) do not model savings decisions of individuals. Accounting for savings behavior in the model, French and Jones (2011) find that individuals with RHI retire half a year earlier than individuals who try to secure health insurance through work.

A few studies use the Current Population Survey (CPS) to investigate the relationship between health insurance and early retirement. Gruber and Madrian (1996) exploit the mandates that allow individuals to purchase group health insurance from their employers. The study uses the Merged Outgoing Rotation Group (MORG) of the CPS for the years between 1980 and 1990. Using a probit model, they find an increase in early retirement by 5.4 percent vis-à-vis an increase in continuation coverage by 1 year. Gruber and Madrian (1995) also find a positive relationship between the probability of retirement and continuation coverage such that the hazard ratio increases by 32.4 percent with 1 year of continuation coverage. Boyle and Lahey (2010) use the health insurance expansion of the U.S. Department of Veterans Affairs as a natural experiment in a difference-in-differences (DD) setup to analyze the labor supply of older veterans. Using the March CPS for the years 1992-2002, the authors show a 3.3 percent decrease in the probability of employment after the expansion. The findings also suggest an 8.4 percent increase in part-time employment.[12]

Nyce et al. (2013) exploit employee-level data obtained from the clients of Tower

[12] The effect on self-employment is negative, but it is not statistically different from zero.

Watson, a benefits consulting firm, for the years 2005 through 2009. The data provide employee records of 54 firms, as well as information on the size of the employer contribution towards health coverage. The findings imply that the probability of not being employed (defined as "turnover") increases by 36 percent at age 62 for workers with subsidized coverage. The increase in the probability of turnover is 49 percent and 38 percent when ages are 63 and 64, respectively. The authors do not find significant effects for those who do not meet the eligibility criteria based on years of service for coverage contribution.

Similar to Shoven and Slavov (2014), Fitzpatrick (2014) analyzes the retirement decision of public sector workers given the availability of RHI. Using administrative data from Illinois Public Schools (IPS) for the school years 1970-1971 and 1991-1992, the author specifically focuses on public school teachers who must have 8 years of tenure to qualify for retirement benefits. The author shows that workers who are eligible for RHI retire 2 years earlier.[13]

All of the studies above show a positive relationship between RHI (or continuation coverage) and early retirement. These studies focus on RHI due to limited alternatives of health insurance that may incentivize early retirement. An alternative health insurance option, however, is made available under the ACA's Medicaid expansion. Although Medicaid is viewed as a substitute for RHI, its effects on retirement need to be formally investigated. To my knowledge, there are five studies investigating the aforementioned relationship, and the findings are mixed. Using the basic monthly CPS, Levy, Buchmueller, and Nikpay (2016) stratify the sample into expansion and non-expansion states, and test for a differential trend in retirement for individuals aged 50-64. The estimates do not yield statistically significant effects on the probability of retirement and part-time employment in the post-2014 period. When all adults are pooled together in the sample, it is likely to underestimate the effect of Medicaid on retirement due to the fact that parents have access to Medicaid in both expansion and non-expansion states in the pre- and post-2014 period.[14]

Gustman, Steinmeier, and Tabatabai (2018) use the HRS to look at the effect of the ACA on retirement for the years 2010-2014. They do not find any impact on retirement

[13] One limitation of the study is with regards to external validity because the author drops public schools in large cities.

[14] As discussed by Aslim (2016), large and heterogeneous treatment groups may jeopardize the estimates for labor market outcomes.

resulting from the ACA. Due to the limitation of their data set with respect to the sample period, their benchmark DD analysis does not incorporate the Medicaid expansion, which is one of the major provisions of the ACA that significantly reduced the uninsured rate. In fact, Frean, Gruber, and Sommers (2017) show that 30 percent of the ACA's impact on coverage gains came after 2014.[15] Depending on the degree of job lock, the ACA provisions between 2010 and 2014 may not be effective in incentivizing a decrease in labor supply.[16] In order to study the retirement effects of the ACA over a longer period, the authors consider expected retirement age as an outcome variable, as well as a simulation analysis based on a structural model. When the outcome variable is expected retirement age, the coefficients are close to zero and also statistically insignificant. The authors provide two explanations for these findings. Both explanations imply that the time period after the ACA's implementation may have been too short to internalize the incentives under the ACA, which also does not allow for reoptimization of retirement expectations. The structural model, on the other hand, yields a 0.1 percentage point reduction in full-time employment.

Similar to Gustman, Steinmeier, and Tabatabai (2018), Ayyagari (2017) uses an expected measure of retirement, referred to as the subjective probability of continuing full-time employment past age 62. Ayyagari (2017) uses more waves of the HRS than Gustman, Steinmeier, and Tabatabai (2018), but the sample period also ends in 2014. The sample is restricted to adults between the ages of 45 to 60 who are working full time in 2008. Using a DD approach, the author finds a 5.6 percentage point decline in the likelihood of continuing work past age 62, which corresponds to a 9.9 percent decline relative to the pre-ACA mean. Contrary to Gustman, Steinmeier, and Tabatabai (2018), the expected retirement age is, on average, lower by 3.6 to 7.2 months. These findings suggest a decrease in job lock among workers who do not have access to retiree benefits.

Baughman (2018) explores the effect of two important provisions of the ACA, Medicaid expansions and the creation of health insurance exchanges, on the retirement decisions of men and women aged 51 to 64. Since it is very difficult to obtain data on premiums in the non-group market for the pre- and post-2014 period, the author

[15] Previous studies also show an evidence of reduced working hours vis-á-vis Medicaid expansions (Garthwaite, Gross, and Notowidigdo, 2014, Aslim, 2016, Dague, DeLeire, and Leininger, 2017).

[16] For adults aged 51 to 56 years, some of the effective provisions of the ACA within those periods are the changes to private health insurance with respect to pre-existing conditions, the introduction of health insurance exchanges, and some early expansions.

exploits the state-level regulations (guaranteed issue, ratings restriction, and policies on pre-existing medical conditions) that affected premiums prior to the ACA. The data come from the March CPS for the years 2011-2016. Using a DD model, the overall sample yields no effect on retirement with respect to Medicaid expansions. When the sample is restricted to adults aged 60-64 or low-educated adults, defined as having less than a high school education, there is a statistically significant increase in the probability of leaving the labor force by 2.8 percentage points and 13.2 percentage points, respectively. In addition, the effect of the non-group market regulations on retirement is also statistically significant for targeted populations. For example, interacting the policy variables with low-educated adults or those who are closer to the Medicare eligibility cutoff (adults aged 60-64) yields a positive effect on the probability of leaving the labor force for at least one non-group market regulation.

This study distinguishes from previous studies, especially those that look at the retirement effects of the ACA, in the following aspects. First, I specifically focus on the ACA's Medicaid expansion, one of the largest provisions of the ACA in terms of coverage gains, and I restrict the sample to low-educated childless adults – those who are most likely affected by the expansions. It is important to note that including parents in the analysis may confound the estimates on labor supply due to "woodwork effects" (Sonier, Boudreaux, and Blewett, 2013, Aslim, 2018). Second, most of the previous studies do not use insurance coverage as an independent variable due to endogeneity concerns, whereas I consider actual Medicaid enrollment, and alleviate these concerns using an instrumental variables (IV) model. Note that omitting key variables that capture job characteristics and health status also bias the estimates.[17] For example, Gruber and Madrian (1996) include state fixed-effects to reduce any bias that may result from omitting job characteristics, as the decision to implement mandates may be correlated with job characteristics. In the benchmark model, I control for health, class of workers, and occupation.

[17]In addition, some of the studies that look at the relationship between RHI and retirement have confounded the estimates by not controlling for defined benefit pensions. Note that defined pension plans, which are correlated with the availability of RHI, may increase the probability of early retirement (Gustman and Steinmeier, 1994).

3. Theoretical Effects of Medicaid Enrollment on Retirement

This section introduces a simple static model for a representative household to investigate the relationship between early retirement, modeled as hours of leisure (L), and health insurance (I) that captures the availability of Medicaid. In order to construct the basis of the model, I follow French and Jones (2011) closely. The model, however, differentiates with respect to the production of health, health insurance, and household characteristics. French and Jones (2011) include households above the age of 64 to show the impact of Medicare on labor supply.[18] Here, the main group of interest is adults between the ages of 55 and 64 who are qualified for Medicaid under the Affordable Care Act.

The objective of a representative household is to maximize utility, which consists of consumption, C, and leisure, L,

$$U(C, L), \tag{1}$$

where U is strictly concave in both goods. The first constraint faced by the individual is a time constraint:

$$T = L + N + H, \tag{2}$$

where T is the total time available and N is the hours allotted to work. It is reasonable to exclude fixed costs resulting from employment and labor market reentry because having both decisions exogenous does not change the outcome of the analysis. The loss of leisure due to time spent sick is captured by H. The health production function depends on health insurance (I) via access to medical care and all other factors (X) including gender, age, and education. This allows us to define sick days as $H = H(I, X)$. The second constraint in the model is a budget constraint:

$$Y = C + M, \tag{3}$$

where household income, Y, is a function of all government transfers/benefits (Social Security, financial aid etc.), fringe benefits including pensions, spouse's income, asset income, wages, and hours worked. Without loss of generality, I assume that wage (w)

[18] Using a dynamic model, they also simulate the effect of employer-provided health insurance on labor force participation rates.

is fixed, and does not vary with working hours and health. The household income is defined as Y, where $Y = w(T - L - H) + A$. Note that making Y a function of all taxable income/benefits (A) that do not vary with L does not affect the equilibrium choice.

Medical expenses depend on the health production function and X: $M = M(H, X)$. I assume that M is strictly concave in H (and C is also strictly concave in M).[19] In terms of modeling health insurance (I), I consider the case where I is continuous and exogenous, and the case where I is defined as a probability function conditional on leisure, $I = Prob(\text{Medicaid} = 1|L) = \Phi(L)$, which takes into account the Medicaid notch. The aim here is to capture a positive relationship between leisure and the likelihood of having Medicaid. Since the main focus of the study is on adults who qualify for Medicaid, I do not exploit the availability of RHI. In addition, most of the low-income adults have limited access to retiree benefits offered by full-time jobs.[20]

If I is exogenously determined, this eliminates the issue of self-selecting into Medicaid by reducing working hours. The effect of I on retirement (or optimal leisure), L^*, could be unambiguously determined from the constraints. It is implicit by the constraints (2) and (3) that $\frac{\partial H}{\partial I} < 0$ and $\frac{\partial M}{\partial H}\frac{\partial H}{\partial I} < 0$ not only lead to more time spent on leisure and work, but also generate a positive income effect on older adults. Assuming C and L are normal goods, this would imply an increase in both goods with respect to Medicaid. The comparative static obtained from the unconstrained optimization also supports this observation:

$$\frac{\partial L^*}{\partial I} = -\frac{\frac{\partial MU_C}{\partial C}\frac{\partial^2 C}{\partial M^2}\frac{\partial^2 M}{\partial H^2}\frac{\partial H}{\partial I}}{\frac{\partial MU_L}{\partial L}}, \tag{4}$$

where MU_C and MU_L are marginal utilities with respect to C and L, respectively. The signs of the partial derivatives are:

$$\underbrace{\overbrace{\frac{\partial MU_C}{\partial C}}^{<0}\overbrace{\frac{\partial^2 C}{\partial M^2}\frac{\partial^2 M}{\partial H^2}}^{>0}\overbrace{\frac{\partial H}{\partial I}}^{<0}}_{>0} \text{ and } \underbrace{\frac{\partial MU_L}{\partial L}}_{<0} \Rightarrow \frac{\partial L^*}{\partial I} > 0. \tag{5}$$

An important concern that needs to be addressed is self-selection into Medicaid by

[19]Since marginal medical care received decreases as health worsens, it is reasonable that M is nonlinear in H.

[20]In a dynamic framework, however, it would be interesting to investigate the employment outcomes of individuals who choose between RHI and Medicaid.

manipulating income through working hours. As also defined by Yelowitz (1995), there is a break even point for leisure (or working hours), often referred to as the Medicaid notch, such that any leisure beyond that point would allow an individual to qualify for Medicaid. Hence, this may give an incentive to individuals, especially those with poor health, to reduce working hours. For this case, Medicaid could be modeled as a function of L, where the probability of qualifying for Medicaid increases as L increases. This case, however, does not contradict to I being exogenous because the time and income effect still work in the same direction. This implies that, accounting for self-selection, an individual could improve health outcomes (less sick days) and reduce medical expenses through Medicaid, which then results in more income and total time to spend on leisure. Note that the model does not capture any savings decisions and/or family considerations that may give more insights about the retirement decision vis-à-vis the changes in income. Taking home production and/or within-household division of labor into account, women (especially wives) may change their labor market behavior depending on the labor supply of men. For example, Boyle and Lahey (2016) explore the labor supply of wives of older male veterans before and after the Veterans Affairs (VA) health benefits expansion. The authors find an increase in the labor supply of wives as husbands' labor supply decreases, suggesting that financial incentives dominate leisure complementarity among couples. I test for possible heterogeneity in retirement decisions of men and women in the empirical analysis.

4. Selection of Expansion and Non-Expansion States

The timing of expansions is crucial in order to disentangle the causal effect of Medicaid on early retirement. Although most expansion states raised their income eligibility limits to 138% FPL in January 2014, some expanded coverage for childless adults earlier than 2014 – referred to as early expansion states. In addition, a number of states expanded coverage after January 2014, and I refer to those as late expansion states. Early expansion states differ not only in their timing of expansions, but also with respect to the coverage benefits provided for childless adults. Column (3) in Table 1, on one hand, includes the list of early expansion states that provided full coverage for eligible childless adults. Column (1), on the other hand, lists the states that provided limited coverage before the ACA's

Medicaid expansion. The mandatory benefits of Medicaid include, but are not limited to, inpatient and outpatient hospital services, nursing facility services, and laboratory and X-ray services.[21]

There are 13 states that provided limited coverage for adults, mainly access to primary care services, before the expansion in 2014 (see Table 1 for a complete list of states). California, an expansion state, provided limited coverage for adults under the Medicaid Coverage Expansion (MCE) and the Health Care Coverage Initiative (HCCI) before January 2014 (Alker et al., 2013). Utah, a non-expansion state, signed the 1115 Primary Care Network (PCN) Demonstration Waiver in December 2011 that provided limited coverage of primary care services for childless adults. Ten of these states with limited benefits fully expanded Medicaid in 2014 (labeled as "E" in column (1)) or after 2014 (labeled as "LE" in column (1)), and the remaining three states opted-out of the ACA's Medicaid expansion in 2014 (labeled as "NE" in column (1)).

In column (2) of Table 1, there are two states with closed enrollment that provided full coverage for eligible adults before January 2014. In 2000, Arizona expanded Medicaid coverage for childless adults below 100% FPL. In May 2011, over 200,000 childless adults enrolled in the program. On July 8, 2011, Arizona decided to freeze enrollment in order to redesign its Medicaid program to reduce costs. The Centers for Medicare and Medicaid Services (CMS) approved Arizona's new Section 1115 waiver on October 21, 2011. In 2014, Arizona expanded Medicaid for childless adults below 138% FPL. Colorado, on the other hand, provided Medicaid coverage to "jobless" childless adults below 10% FPL in May 2012. The state capped the enrollment to 10,000 adults. Similar to the case in Arizona, Colorado fully expanded Medicaid to 138% FPL in 2014.

The main analysis includes expansion (E) and non-expansion states (NE) in columns (1) and (2) because the effects on enrollment and/or benefits are limited. The early expansion states in column (3) and late expansion states in column (4) are excluded from the analysis to capture the main effect of the ACA's Medicaid expansion in January 2014. I probe the robustness of the estimates to the inclusion of early and late expansion states. A consistent selection of states in the main analysis is very important for the studies on the ACA's Medicaid expansion. Levy, Buchmueller, and Nikpay (2016), for example, exclude California, Massachusetts, and Arizona (a closed enrollment state) in the main

[21] See full list of mandatory benefits at `http://www.Medicaid.gov`.

analysis by assuming that three states have provided full benefits before January 2014. Note that California and Massachusetts are states with limited benefits. On the other hand, they include Colorado (a closed enrollment state) and the remaining limited benefit states in the main analysis, which contradicts the initial exclusion of three states.

Table 1. State Medicaid Expansion Profile for Childless Adults

(1) States with Limited Benefits (Before January 2014)	(2) States with Closed Enrollment (Before January 2014)	(3) Early Expansion States (Before January 2014)	(4) Expansion States (E) (January 2014)	(5) Late Expansion States (LE) (After January 2014)	(6) Non-Expansion States (NE) (As of December 2016)
California (E)	Arizona (E)	Connecticut	Arizona	Alaska	Alabama
Iowa (E)	Colorado (E)	Delaware	Arkansas	Indiana	Florida
Maine (NE)		Hawaii	California	Louisiana	Georgia
Maryland (E)		Minnesota	Colorado	Michigan	Idaho
Massachusetts (E)		New York	Illinois	Montana	Kansas
Michigan (LE)		Vermont	Iowa	New Hampshire	Maine
New Jersey (E)		District of Columbia	Kentucky	Pennsylvania	Mississippi
New Mexico (E)			Maryland		Missouri
Oklahoma (NE)			Massachusetts		Nebraska
Oregon (E)			Nevada		North Carolina
Utah (NE)			New Jersey		Oklahoma
Washington (E)			New Mexico		South Carolina
Wisconsin* (E)			North Dakota		South Dakota
			Ohio		Tennessee
			Oregon		Texas
			Rhode Island		Utah
			Washington		Virginia
			West Virginia		Wyoming
			Wisconsin*		
$n = 13$ states	$n = 2$ states	$n = 7$ states	$n = 19$ states	$n = 7$ states	$n = 18$ states

Notes: Columns (3), (4), (5), and (6) are mutually exclusive (51 states in total). (E) indicates an expansion state, (NE) indicates a non-expansion state, (LE) indicates a late expansion state. The main analysis includes (E) and (NE) states in columns (1) and (2) because of limited benefits compared to full Medicaid and/or limited enrollment. *Although Wisconsin opted-out of the ACA's Medicaid expansion, childless adults below 100% FPL are eligible for Medicaid. In the analysis, Wisconsin is treated as an expansion state due to the high eligibility limit. Arizona closed enrollment on July 8, 2011. Colorado closed enrollment on May 1, 2012.
Source: Kaiser Family Foundation, `https://www.kff.org/medicaid/report/annual-updates-on-eligibility-rules-enrollment-and/`, retrieved February 21, 2018.

5. Data

I use data from the Public Use Micro Samples (PUMS) of the American Community Survey (ACS) for the years 2010 through 2016. The ACS provides information on health insurance, health[22], labor market outcomes, and demographic characteristics for the pre- and post-ACA period. To my knowledge, this is the first study to use the ACS to investigate the effect of the ACA's Medicaid expansion on retirement. One of the benefits of using the ACS is having access to geographic identifiers, which allows me to test for the differences in outcomes with respect to states.[23] For this study, I exploit the state-level variation in Medicaid expansions to construct an instrument for Medicaid enrollment. There is also a relatively large sample of older adults in the ACS. However, it is not possible to have a dynamic retirement framework since the ACS does not track the same individuals over time.[24] The Survey of Income and Program Participation (SIPP) could be an alternative data set for retirement studies, but the latest release of the data is for 2014 as of the writing of this paper.

The population that would most likely be affected by the ACA's Medicaid expansion is low-income childless adults. Since education is strongly correlated with income, I restrict the sample to low-educated adults without children under the age of 18. I define low education as having a high school education or less.[25] This restriction reduces concerns of isolating the effects of Medicaid expansions from the effects of simultaneously implemented health insurance exchanges. Moreover, I limit the sample to adults between the ages of 55 and 64.[26] Previous studies show heterogeneous effects on the labor market behavior of men and women (Kapur and Rogowski, 2011, Boyle and Lahey, 2016, Heim and Lin, 2017). Thus, I stratify the sample by gender in all specifications.

The outcome variable, *Retirement*, is an indicator variable taking the value 1 if a person has retirement income in the past 12 months and 0 otherwise.[27] *Medicaid*, on the

[22] Health variables are mainly related to disabilities, including self-care difficulty, hearing difficulty, vision difficulty, independent living difficulty, ambulatory difficulty, and cognitive difficulty.

[23] Note that the publicly available HRS does not include geographic identifiers.

[24] A minor limitation is the absence of survey months that could be used to capture the monthly variation in policy variables. Note that March CPS does not also vary by months.

[25] This restriction is consistent with the literature; for example, Kaestner et al. (2017) restrict the sample to low-educated adults to explore the effect of Medicaid expansions on labor supply.

[26] Note that the availability of Medicare may confound the estimates on Medicaid. Thus, I exclude adults above the age of 64. In addition, some studies denote stronger retirement incentives for adults who are closer to the 64 age cutoff; I also test this by restricting the age to 59-64.

[27] Note that retirement income is highly correlated with the probability of leaving the labor force. An

other hand, shows whether an individual is enrolled for Medicaid coverage or not. For both men and women, the probability of enrolling for Medicaid is higher in expansion states than non-expansion states. Figure 1 shows that Medicaid enrollment among older adults is fairly stable (and close to zero) in non-expansion states. In expansion states, however, there is a spike in Medicaid enrollment after the expansions in 2014. Although women have a higher probability of retirement in non-expansion states, the gap closes in 2014 with an increase in the probability of retirement in expansion states and a decrease in non-expansion states. The trend in retirement is parallel for men in the pre- and post-2014 period. These trends imply that men and women have different labor market responses vis-à-vis health insurance enrollment, which in fact support the previous findings in the literature. Additionally, older women are more likely to work for non-for-profit businesses or the government, whereas older men are more likely to work for for-profit businesses or choose to be self-employed (see Table 2). This also gives additional insights on why labor market behavior differs for men and women.

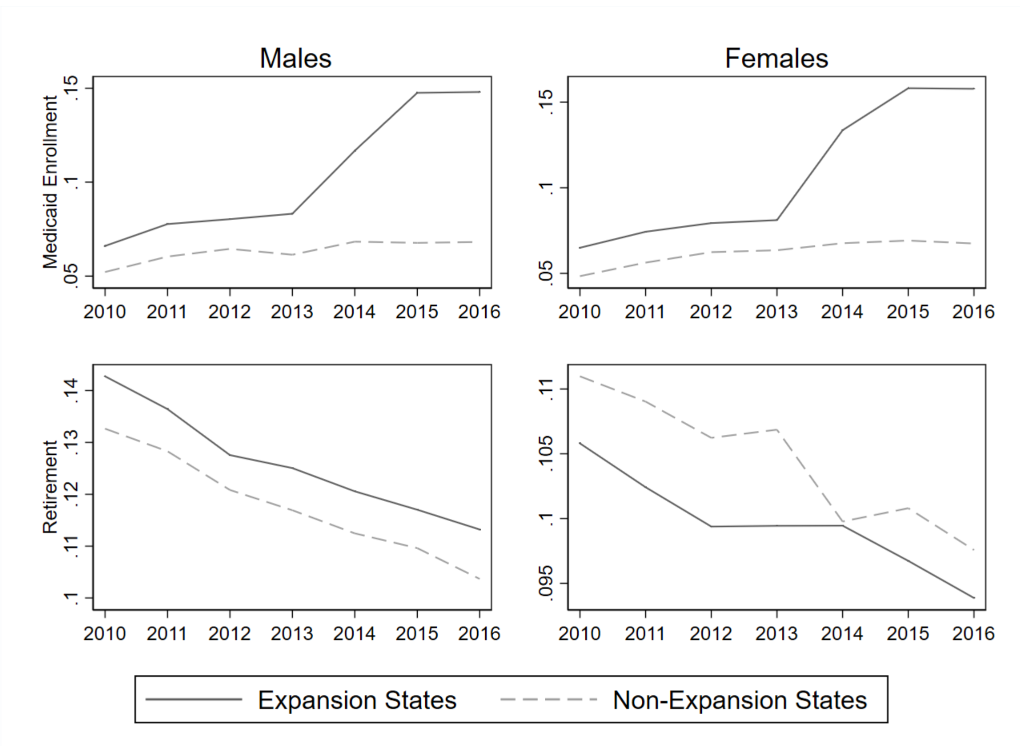

Figure 1. Trends in Medicaid Enrollment and Retirement by State and Gender, ACS 2010-2016

alternative definition could be the probability of leaving the labor force conditional on working full-time in the past 12 months (see, for example, Heim and Lin, 2017). The findings of the IV model are robust to the changes in the definition of the outcome variable and are available upon request.

Table 2. Descriptive Characteristics of Low-Educated Childless Adults Aged 55-64, ACS 2010-2016

	Men			Women		
	All	Expansion States	Non-Expansion States	All	Expansion States	Non-Expansion States
Retirement	0.109	0.113	0.104	0.095	0.093	0.097
Medicaid	0.084	0.106	0.060	0.087	0.111	0.060
Age	58.868	58.871	58.865	59.162	59.149	59.177
White	0.768	0.773	0.762	0.758	0.765	0.750
Black	0.115	0.072	0.165	0.127	0.081	0.179
Asian	0.039	0.055	0.021	0.051	0.069	0.029
Hispanic	0.195	0.217	0.170	0.163	0.182	0.141
Married	0.679	0.685	0.673	0.589	0.596	0.580
Widowed	0.028	0.026	0.030	0.091	0.086	0.096
Divorced	0.171	0.163	0.180	0.207	0.200	0.214
Separated	0.028	0.025	0.031	0.033	0.031	0.035
Less than middle school	0.145	0.155	0.133	0.109	0.123	0.093
High school dropout	0.173	0.156	0.192	0.142	0.130	0.157
High school diploma	0.683	0.689	0.676	0.749	0.748	0.751
Disability	0.165	0.155	0.177	0.147	0.138	0.157
For-profit business	0.698	0.705	0.690	0.673	0.672	0.675
Non-for-profit business	0.033	0.036	0.030	0.072	0.076	0.067
Working for the government	0.099	0.096	0.102	0.146	0.145	0.149
Self-employed	0.160	0.153	0.167	0.097	0.096	0.098
N	302,998	161,929	141,069	272,305	144,401	127,904

Notes: Early and late expansion states are excluded from the analysis (see Table 1). ACS individual-level weights are used in computing means.

6. Methods

The econometric model for the relationship between health insurance enrollment and retirement can be written as:

$$y_{ist} = \beta_0 + \beta_1 Medicaid_{ist} + X'_{ist}\beta_2 + \delta_{1t} + \sigma_{1s} + [\zeta_{1st}] + \epsilon_{ist}. \tag{6}$$

The dependent variable, y_{ist}, is an indicator variable on retirement for individual i in state s at time t (year). $Medicaid$ takes the value 1 if an individual is enrolled for Medicaid and 0 otherwise. The X_{ist} vector includes the following characteristics that vary at the individual-state-year level: age, age-squared, sex, marital status (married, widowed, divorced, separated), race (white, black, Asian), ethnicity (Hispanic), education (less than middle school, high school dropout), indicator variables for disability, class of worker (working for a private for-profit business, working for a private not-for-profit organization, working for the government, or self-employed), and occupation. Additional controls include year fixed effects (δ_{1t}), state fixed effects (σ_{1s}), and state time-varying effects (ζ_{1st}), which are state unemployment rates. Standard errors are clustered at the state-level to account for any serial correlation within states of similar characteristics. Estimates are weighted using the appropriate individual-level weights in the ACS.

As the theory suggests in this paper, there is a discontinuous change in benefits with respect to working hours – the "Medicaid-notch." This specification, however, does not account for self-selection into Medicaid by reducing working hours. To disentangle the causal effect of health insurance enrollment on retirement, I exploit the expansion decision and timing of Medicaid expansions using a difference-in-differences (DD) model (see Table 3). First, I estimate the relationship between enrollment and Medicaid expansions using the following regression:

$$Medicaid_{ist} = \alpha_0 + \alpha_1 Expansion_s * Post_t + X'_{ist}\alpha_2 + \delta_{2t} + \sigma_{2s} + [\zeta_{2st}] + \epsilon_{ist}, \tag{7}$$

where $Expansion$ is an indicator variable on whether a state is an expansion state or not, and $Post$ is an indicator variable that takes the value 1 in the post-2014 period and 0 otherwise.[28] Identifying α_1 as the casual effect of Medicaid expansions on enrollment rests on the standard DD assumption of parallel trends. Second, I estimate the relationship between retirement and Medicaid expansions using the following regression:

$$y_{ist} = \gamma_0 + \gamma_1 Expansion_s * Post_t + X_{ist}^{'}\gamma_2 + \delta_{3t} + \sigma_{3s} + [\zeta_{3st}] + \epsilon_{ist}. \tag{8}$$

The standard parallel trends assumption is also required to interpret γ_1 as the casual effect of Medicaid expansions on retirement. For specifications (7) and (8), I test the plausibility of this assumption by including a set of interactions of the expansion states with years, $\sum_{t=2011}^{2013} Expansion_s * Year_t$, where $Year$ refers to year fixed effects with 2010 being the base year. As an additional check, I replace $Expansion_s * Post_t$ with $\sum_{t=2011}^{2016} Expansion_s * Year_t$. If the estimates prior to 2014 are statistically insignificant, this would imply that Medicaid enrollment and retirement do not trend differentially in the absence of Medicaid expansions. I also restrict the pre-policy period to 2011-2013 to check whether the benchmark estimates are robust to potential identification threats that may result from the Great Recession.[29]

Taking the ratio of two intention-to-treat (ITT) effects, γ_1/α_1, yields the treatment-on-the-treated (TOT). This is equivalent to an IV analysis, where the IV effect is a Wald estimate given a just-identified system. The only channel by which expansion states can affect retirement is through Medicaid enrollment, and hence the exclusion restriction is satisfied. Since the purpose of the ACA's Medicaid expansion is to reduce the rate of uninsured, it is reasonable to assume that the policy is exogenous with respect to retirement. In addition to exclusion restriction and policy exogeneity, I assume that there are no spillovers resulting from the treatment that changes the outcome of other childless adults, which is referred to as the stable unit treatment value assumption (SUTVA).

[28] When early and late expansion states are included in the analysis, the timing of expansion changes, and *Post* also changes accordingly.

[29] The findings for the 2009-2013 pre-policy period are in the Appendix.

7. Results

7.1. Main Results

Table 3 presents the results from estimating the DD regression in Equation (7). In columns (4) and (8), men increase Medicaid enrollment by 5 percentage points after the Medicaid expansions. For women, enrollment in Medicaid increases by 6 percentage points. Note that the Medicaid expansions reduce private insurance enrollment by 3.1 percentage points for men and 4.3 percentage points for women, suggesting a crowd-out rate of 62 percent and 72 percent, respectively.[30] Although I focus on the most inclusive regression in columns (4) and (8), the estimates are robust across specifications with control variables. Additionally, the estimates are not sensitive to restricting the pre-policy period to 2011-2013 and 2009-2013 (see Appendix, Table A1). These findings are similar in magnitude to those reported in existing studies. Kaestner et al. (2017), for example, show an increase in Medicaid enrollment among childless adults by 4 percentage points using the 2010-2014 ACS. Using the same data, Leung and Mas (2016) find the change in Medicaid coverage to be 3 percentage points among childless adults.

Table 4 contains the DD estimates for retirement obtained from Equation (8). For men, the intent-to-treat (ITT) effects are of opposite signs and not statistically different from zero. This finding is consistent across samples and specifications. Women, however, increase their probability of retirement by around 0.6 percentage points in columns (4) and (8). This finding supports studies that show different retirement behaviors for men and women given health insurance expansions (see, for example, Boyle and Lahey, 2016, Heim and Lin, 2017). As discussed in the theory section, women may change their retirement behavior given the labor supply of men. The findings here suggest that men do not retire early, which may incentivize women to reduce labor supply considering within-household division of labor and lower opportunity costs.

The validity of DD estimates reported in Tables 3 and 4 hinges on the parallel trends assumption. To investigate the plausibility of this assumption, I estimate event study specifications for Medicaid enrollment and retirement (see Table 5). Since retirement effects are only observed for women, I do not report the estimates for men. The regression

[30] The estimates for private health insurance are available upon request.

includes state and year indicators, demographic and employment characteristics, and state time-varying effects. In all specifications and samples, there are no statistically significant differences between the treatment and control groups before 2014. In the post-ACA period, the increase in Medicaid enrollment is around 6 percentage points (see columns (1) and (5)), and the increase in the probability of retirement is in between 0.5 and 0.6 percentage points (see columns (3) and (7)). Although Medicaid enrollment increases in the post-ACA period, low-educated women significantly increase their probability of retirement in 2014 (see columns (4) and (8)). The increase in outreach prior to the ACA's Medicaid expansion might have reduced uncertainty regarding the timing of retirement (Ayyagari, 2017).

Since the parallel trends assumption is supported by the estimates from the event study specifications, I turn to Wald estimates by taking the ratio of the two intention-to-treat (ITT) effects reported in Tables 3 and 4. The ratios for men and women are presented in Table 6. For men, I do not find any significant effect of Medicaid enrollment on early retirement. For women, Medicaid enrollment increases the probability of retirement by 10.1 ($\gamma_1/\alpha_1 = 0.0061/0.0605$) and 10.3 ($\gamma_1/\alpha_1 = 0.00627/0.0607$) percentage points in columns (2) and (4), respectively. Those who are full-time workers are more likely to experience job lock due to higher probability of having access to ESHI. If this effect for women is not spurious, then Medicaid enrollment should have a larger effect on early retirement for those who are employed full-time in the past 12 months. In fact, Table A2 (see Appendix) shows that women with Medicaid increase their probability of retirement by 12 to 13 percentage points. These estimates are larger than the OLS estimates, which are around 1 percentage point.[31]

To test whether the IV and OLS estimates are different, I perform a Durbin-Wu-Hausman test for the 2010-2016 sample period. The F-statistic is 10.08 (P-value = 0.003), which implies that I can reject the null hypothesis that Medicaid enrollment is exogenous. Thus, there are two plausible explanations for large IV estimates: i) choice of instrument and ii) measurement error in Medicaid enrollment. The first stage is relatively good strong, as shown by F-statistic over 50 on the excluded instrument. Since expansion states can affect labor supply only through actual enrollment, the exclusion restriction on the expansion status is also satisfied. These

[31]The OLS estimates are available upon request.

alleviate concerns regarding the validity of the instrument. As shown in the calculation of Wald estimates, it is the first stage that inflates the estimates. The first stage estimates, however, are reasonable given that the population analyzed is low-educated childless adults who are most likely affected by the Medicaid expansions. As discussed above, the first stage estimates are similar to those found in the literature. On the other hand, measurement error in public health insurance reporting in the ACS could bias the IV estimates (Boudreaux et al., 2015). Note that the IV will yield larger estimates than the OLS when the measurement error is random. Even if this is the case, these findings are still informative in providing an upper bound for the effects of Medicaid expansion on early retirement.

Table 3. Effect of the ACA's Medicaid Expansion on Medicaid Enrollment: DD Estimates for Low-Educated Men and Women

Sample period	2010 - 2016				2011 - 2016			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Men								
Expansion*Post	0.0730***	0.0541***	0.0543***	0.0500***	0.0720***	0.0540***	0.0539***	0.0491***
	(0.01611)	(0.01235)	(0.01237)	(0.00848)	(0.01644)	(0.01265)	(0.01270)	(0.00869)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effect	No	No	No	Yes	No	No	No	Yes
N	302,998	302,998	302,998	302,998	264,412	264,412	264,412	264,412
B. Women								
Expansion*Post	0.0901***	0.0646***	0.0647***	0.0605***	0.0887***	0.0654***	0.0657***	0.0607***
	(0.01783)	(0.01228)	(0.01219)	(0.00828)	(0.01792)	(0.01203)	(0.01196)	(0.00796)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	272,305	272,305	272,305	272,305	234,752	234,752	234,752	234,752

Notes: Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table 4. Effect of the ACA's Medicaid Expansion on Retirement: DD Estimates for Low-Educated Men and Women

Sample period	2010 - 2016				2011 - 2016			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Men								
Expansion*Post	-0.00527	-0.00118	-0.00104	-0.00225	-0.00241	-0.00089	-0.00068	-0.00165
	(0.00885)	(0.00540)	(0.00520)	(0.00505)	(0.00918)	(0.00500)	(0.00489)	(0.00485)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effect	No	No	No	Yes	No	No	No	Yes
N	302,998	302,998	302,998	302,998	264,412	264,412	264,412	264,412
B. Women								
Expansion*Post	-0.00628	0.00688***	0.00605***	0.00610***	-0.00527	0.00700***	0.00608***	0.00627***
	(0.00589)	(0.00206)	(0.00180)	(0.00182)	(0.00613)	(0.00208)	(0.00185)	(0.00189)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	272,305	272,305	272,305	272,305	234,752	234,752	234,752	234,752

Notes: Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table 5. Effect of the ACA's Medicaid Expansion on Medicaid Enrollment and Retirement: Event Study Estimates for Low-Educated Women

Sample period	2010 - 2016				2011 - 2016			
Dependent variable	Medicaid		Retirement		Medicaid		Retirement	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Expansion*Post	0.0587***		0.00612**		0.0636***		0.00520**	
	(0.00984)		(0.00274)		(0.00633)		(0.00244)	
d_{2011}*Expansion	-0.00514	-0.00514	0.00104	0.00104				
	(0.00572)	(0.00566)	(0.00260)	(0.00260)				
d_{2012}*Expansion	0.00215	0.00204	0.00099	0.00101	0.00754	0.00742	-0.00011	-0.00008
	(0.00593)	(0.00587)	(0.00339)	(0.00339)	(0.00557)	(0.00558)	(0.00380)	(0.00380)
d_{2013}*Expansion	-0.00406	-0.00419	-0.00199	-0.00197	0.00133	0.00118	-0.00313	-0.00310
	(0.00668)	(0.00658)	(0.00497)	(0.00495)	(0.00723)	(0.00715)	(0.00467)	(0.00465)
d_{2014}*Expansion		0.0446***		0.00899**		0.0497***		0.00798**
		(0.00886)		(0.00373)		(0.00539)		(0.00320)
d_{2015}*Expansion		0.0647***		0.00468		0.0697***		0.00376
		(0.01286)		(0.00366)		(0.00960)		(0.00342)
d_{2016}*Expansion		0.0671***		0.00461		0.0719***		0.00375
		(0.01029)		(0.00432)		(0.00769)		(0.00439)
N	272,305	272,305	272,305	272,305	234,752	234,752	234,752	234,752

Notes: All specifications include state and year fixed effects, demographic and employment characteristics, and state time-varying effects. Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table 6. Effect of Medicaid Enrollment on Retirement: Wald Estimates for Low-Educated Men and Women

Sample period	2010 - 2016		2011 - 2016	
Subgroup	Men	Women	Men	Women
	(1)	(2)	(3)	(4)
Medicaid	-0.0406	0.1011***	-0.0336	0.1033***
	(0.10299)	(0.03042)	(0.10062)	(0.03091)
F-statistics on excluded instrument	34.23	52.49	31.87	58.12
N	302,998	272,305	264,412	234,752

Notes: All specifications include state and year fixed effects, demographic and employment characteristics, and state time-varying effects. Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

7.2. Robustness and Falsification Checks

Since the benchmark analyses exclude early expansion states (7 states) and late expansion states (7 states), I test whether the DD estimates on Medicaid are robust to the inclusion of these 14 states (see Table 7). I use 2011-2016 as the sample period since the estimates are less likely to be confounded by the Great Recession.[32] For early and late expansion states, the indicator variable *Post* changes with the timing of expansions instead of 2014. In column (4), Medicaid enrollment increases for men and women by 3.6 percentage points and 4.3 percentage points, respectively. The ACA's Medicaid expansion has no effect on the retirement decision of men, whereas women are more likely to retire early by 0.5 percentage points (see column (8)). Note that including early and late expansion states in the sample may bias the estimates due to woodwork effects. On the one hand, an increase in outreach may lead to higher take up rates in 2014 for early expansion states. On the other hand, as outreach decreases over time there could be less enrollment in late expansion states. Thus, woodwork effects may cause an underestimation of the effect of Medicaid on retirement. The findings in Table 7 are consistent with a priori expectations.

Next, I test the robustness of the estimates in Tables 3 and 4 when the ages of individuals are restricted to 59-64.[33] The estimates are fairly robust compared to the benchmark case (see Table 8). For men and women, Medicaid enrollment increases by 5 percentage points and 6 percentage points, respectively. I still do not find any effect on retirement for men. On the other hand, women increase their probability of retirement by 1.2 percentage points after the expansions. This coefficient is larger compared to 0.6 percentage point in Table 4. This finding implies that Medicaid is more effective on relatively older women who are closer to the full-benefit retirement age.

As a falsification check, I restrict the sample to high-educated adults.[34] Since the ACA's Medicaid expansion mainly targeted low-income adults, there should be no effect (or a limited effect) on the retirement behavior of high-educated men and women. Table 9 contains the DD estimates for high-educated men and women. Compared to the estimates

[32]The findings are, however, robust to changes in the sample period to 2010-2016.

[33]The benchmark sample includes childless adults aged 55 to 64 years. This age restriction for the lower boundary is the same as the previous studies that investigate the effect of RHI on early retirement (Gruber and Madrian, 1996, Rogowski and Karoly, 2000, Boyle and Lahey, 2010, Shoven and Slavov, 2014, Fitzpatrick, 2014). On the other hand, there are some retirement studies that use either 50 or 51 as the lower boundary for age (Strumpf, 2010, Robinson and Clark, 2010, Levy, Buchmueller, and Nikpay, 2016).

[34]Note that these individuals do not have dependent children under the age of 18.

in Table 3 for the years 2011-2016, the increase in Medicaid enrollment is smaller by around 50 percent for both men and women. As predicted, I also do not find any evidence of an increase in early retirement across specifications. Even when the model does not include any control variables and state and year fixed effects, there is a decrease in the probability of retirement for high-educated adults. Overall, this finding justifies the fact that targeted subgroups are relatively more affected by the ACA's Medicaid expansion.

Table 7. Robustness Check: Including Early and Late Expansion States, ACS 2011-2016

Dependent variable	Medicaid				Retirement			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Men								
Expansion*Post	0.0673***	0.0391***	0.0389***	0.0364***	-0.00474	-0.00276	-0.00252	-0.00286
	(0.01084)	(0.00997)	(0.01001)	(0.00737)	(0.00706)	(0.00372)	(0.00366)	(0.00350)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	354,447	354,447	354,447	354,447	354,447	354,447	354,447	354,447
B. Women								
Expansion*Post	0.0794***	0.0451***	0.0452***	0.0428***	-0.00572	0.00571***	0.00522***	0.00529***
	(0.01233)	(0.01037)	(0.01034)	(0.00771)	(0.00418)	(0.00184)	(0.00164)	(0.00158)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	313,859	313,859	313,859	313,859	313,859	313,859	313,859	313,859

Notes: Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table 8. Robustness Check: Age Group 59-64, ACS 2011-2016

Dependent variable	Medicaid				Retirement			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Men								
Expansion*Post	0.0703***	0.0513***	0.0515***	0.0465***	-0.00261	-0.00139	-0.00122	-0.00271
	(0.01741)	(0.01316)	(0.01322)	(0.00935)	(0.01335)	(0.00618)	(0.00590)	(0.00579)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	137,661	137,661	137,661	137,661	137,661	137,661	137,661	137,661
B. Women								
Expansion*Post	0.0872***	0.0648***	0.0645***	0.0593***	-0.00543	0.01118***	0.01076***	0.01151***
	(0.01857)	(0.01227)	(0.01214)	(0.00822)	(0.00886)	(0.00351)	(0.00326)	(0.00321)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	132,495	132,495	132,495	132,495	132,495	132,495	132,495	132,495

Notes: Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table 9. Falsification Check: High-Educated Men and Women, ACS 2011-2016

Dependent variable	Medicaid				Retirement			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
A. Men								
Expansion*Post	0.0368***	0.0274***	0.0274***	0.0259***	-0.0228***	0.00075	0.00114	0.00022
	(0.00387)	(0.00321)	(0.00322)	(0.00247)	(0.00664)	(0.00366)	(0.00360)	(0.00336)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	553,843	553,843	553,843	553,843	553,843	553,843	553,843	553,843
B. Women								
Expansion*Post	0.0413***	0.0318***	0.0316***	0.0306***	-0.0129**	-0.00274	-0.00342	-0.00354
	(0.00345)	(0.00335)	(0.00342)	(0.00327)	(0.00533)	(0.00317)	(0.00328)	(0.00331)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	540,147	540,147	540,147	540,147	540,147	540,147	540,147	540,147

Notes: Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

8. Discussion

There has been an increasing number of studies investigating the relationship between health insurance and labor supply. Various studies showed evidence on job lock where individuals continue working solely to retain health insurance benefits. Policy makers, however, are responding to the spillover effects of the ACA's Medicaid expansion. Arkansas, for example, adopted work requirements for low-income individuals to be eligible for Medicaid. As of January 2018, additional states are waiting to get approval from the CMS for their work requirement waivers. Although the effect of work requirements on enrollment and labor supply is currently ambiguous, the nature of the Medicaid program is changing with the adoption of these welfare rules. Thus, it has become crucial to formally investigate the spillover effects under the ACA.

In this paper, I explore the relationship between Medicaid enrollment and retirement for men and women aged 55-64. I limit the sample to low-educated childless adults, which is a group that is directly targeted under the ACA's Medicaid expansion. I find that the expansions increase Medicaid enrollment by 5 percentage points and 6 percentage points for men and women, respectively. The estimates suggest that the expansions and Medicaid enrollment have no effect on men's retirement behavior. Women, on the other hand, increase their probability of retirement by around 0.6 percentage points. Wald estimates or the treatment-on-the-treated (TOT), computed by taking the ratio of intention-to-treat (ITT) effects for retirement, imply a 10 percentage point increase in the probability of retirement for women. Although the TOT effects seem relatively large, I do not find any issues regarding the validity of the instrument. If IV estimates are possibly inflated due to a random measurement error in Medicaid enrollment, then these estimates should be considered as an upper bound for the retirement effect.

Pairing the theoretical discussion with the empirical findings yields important welfare implications. The simple static model in the paper shows that Medicaid allows beneficiaries to increase leisure and consumption of non-medical goods and services. Thus, the ACA's Medicaid expansion is welfare-improving for target populations. Early retirements, however, increase the tax burden of the Medicaid program on nonparticipants through implicit taxes. An expected consequence of implicit taxes is a reduction in total output.

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Appendix

Table A1. Robustness Check: Sample Period 2009-2016

Dependent variable	Medicaid				Retirement			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
				A. Men				
Expansion*Post	0.0719***	0.0547***	0.0546***	0.0513***	-0.00810	-0.00157	-0.00132	-0.00204
	(0.01596)	(0.01219)	(0.01221)	(0.00887)	(0.00862)	(0.00554)	(0.00534)	(0.00514)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effect	No	No	No	Yes	No	No	No	Yes
N	340,412	340,412	340,412	340,412	264,412	264,412	264,412	264,412
				B. Women				
Expansion*Post	0.0889***	0.0659***	0.0660***	0.0627***	-0.00740	0.00665***	0.00578**	0.00572**
	(0.01755)	(0.01198)	(0.01182)	(0.00851)	(0.00563)	(0.00233)	(0.00211)	(0.00213)
State indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Year indicators	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Demographic characteristics	No	Yes	Yes	Yes	No	Yes	Yes	Yes
Employment characteristics	No	No	Yes	Yes	No	No	Yes	Yes
State time-varying effects	No	No	No	Yes	No	No	No	Yes
N	309,681	309,681	309,681	309,681	(0.006)	(0.002)	(0.002)	(0.002)

Notes: Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.

Table A2. Effect of Medicaid Enrollment on Retirement: Wald Estimates for Full-Time Employed Men and Women

Sample period	2010 - 2016		2011 - 2016	
Subgroup	Men	Women	Men	Women
	(1)	(2)	(3)	(4)
Medicaid	-0.0397	0.1274**	0.0208	0.1215**
	(0.11264)	(0.05131)	(0.11798)	(0.05125)
F-statistics on excluded instrument	21.91	25.80	22.20	28.33
N	213,465	152,323	187,210	131,732

Notes: All specifications include state and year fixed effects, demographic and employment characteristics, and state time-varying effects. Early and late expansion states are excluded from the analysis (see Table 1). Standard errors clustered by state are in parentheses.*Statistical significance at the 10 percent level. **Statistical significance at the 5 percent level. ***Statistical significance at the 1 percent level.