Health Insurance and Job Mobility: Evidence from the Affordable Care Act

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Abstract

Health insurance can shape workers' job mobility by creating incentives to stay with or leave an employer. This paper estimates the causal effect of the Affordable Care Act's Medicaid expansion on job switching among low-income individuals without children, a group largely ineligible for Medicaid prior to the expansion. Using monthly data from the Current Population Survey with a Difference-in-Differences technique, the estimates show that Medicaid expansion reduced monthly job switching by 0.5 percentage points, representing an 18.3% decline in job mobility. The effect is concentrated among those nearing retirement age (55 to 64). These findings suggest that Medicaid expansion reduced the likelihood that low-income childless adults leave their jobs due to concerns about health insurance.

Keywords: Health insurance; Job mobility; Medicaid expansion; Affordable Care Act

JEL Codes: I13, I18, J22, J63, H75

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1 Introduction

Access to health insurance can shape workers' employment decisions, particularly whether to change jobs. Job switching is a key channel of labor market mobility, allowing workers to find better job matches that raise productivity and earnings. When health insurance is tied to employment, however, workers may hesitate to leave a job for fear of losing coverage or may seek jobs that provide it. Public programs such as Medicaid can reduce this dependence by offering coverage outside the workplace. This paper examines how the Affordable Care Act (ACA) Medicaid expansion affected job switching among workers who became newly eligible when eligibility rules changed

In March 2010, President Barack Obama signed the Affordable Care Act (ACA) into law, marking the most significant expansion of the U.S. healthcare system since the introduction of Medicare in 1965. Before the ACA, many low-income adults, especially those without dependent children, fell into a coverage gap. They earned too much to qualify for Medicaid but not enough to afford private insurance. The ACA aimed to increase access to these individuals by expanding public insurance and making private coverage more affordable..

One of the law's key provisions required states to extend Medicaid eligibility to all individuals with incomes below 138 percent of the federal poverty line, beginning in 2014. In 2012, the U.S. Supreme Court ruled that the federal government could not compel states to expand, making the policy optional and leaving the decision to each state. As a result, some states adopted the expansion while others did not, creating a clear natural experiment that allows for a credible comparison between expansion and non-expansion states.

Because Medicaid had previously covered primarily families with children, the expansion most directly affected low-income, nonelderly adults without dependent children, referred to here as childless adults. Several studies, including Courtemanche et al. (2017), Cawley et al. (2018), Leung and Mas (2018), and Kaestner et al. (2017), show that Medicaid expansion substantially increased insurance coverage among this group. For example, Kaestner et al. (2017) report that coverage for childless adults increased by roughly 50 percent. Other research, such as Wherry and Miller (2016), Frean et al. (2017), and Gooptu et al. (2016), finds similar gains when examining the broader low-income population. These findings confirm that the expansion reached its intended target and significantly reduced uninsurance among adults who previously had little or no access to coverage.

Expanding Medicaid eligibility can influence job mobility through two channels that work in opposite directions. Workers with employer-contingent health insurance may hesitate to change jobs because they would lose their coverage, a behavior known as *job lock*. In contrast, uninsured workers may be more likely to leave their current jobs to find new ones that offer health benefits, which is referred to as *job push* (Anderson 1997; Hamersma and Kim 2009). By providing public coverage independent of employment, Medicaid expansion could reduce both types of behavior. It might lessen job lock by allowing insured workers to move more freely, but it could also reduce job push by decreasing the incentive for uninsured workers to seek jobs that provide health insurance. The overall effect on job mobility is therefore ambiguous and must be determined empirically.

In this study, I use the ACA Medicaid expansion as a quasi-experimental setting to estimate its effect on job switching among newly eligible workers. I focus on low-income, non-disabled, childless adults in states that expanded Medicaid in 2014 compared to those in states that did not expand during the study period. Using monthly data from the Current Population Survey (CPS) from 2008 to 2019, I estimate a Difference-in-Differences (DiD) model to measure how job mobility changed in response to the policy.

To account for states that expanded after 2014, I also estimate the model using the Wooldridge (2021) estimator, which addresses potential bias from staggered treatment adoption. The results from both approaches show that Medicaid expansion reduced the probability of job switching by about 0.5 percentage points, or roughly 18 percent relative to the pre-expansion mean. The findings are consistent with the job-push mechanism, in which workers are less likely to change jobs once public insurance reduces the need to seek employer-based coverage. The effect becomes more pronounced about three years after the expansion. Heterogeneity analysis indicates that the decline in job switching is concentrated among workers nearing retirement (ages 55–64), who may be most concerned about access to healthcare.

This study contributes to the literature on Medicaid expansion and job switching in three main ways. First, it focuses on childless adults who became fully eligible for Medicaid for the first time in 2014. Before the ACA, states that wanted to extend coverage to this group faced substantial administrative barriers such as applying for federal waivers. Because these adults gained eligibility only after 2014, they provide a clean test of how the policy influenced job mobility. Second, I use data that extend five years beyond the initial expansion, which allows for an examination of both short- and medium-term effects. Earlier studies such as Gooptu et al. (2016) and Callison and Sicilian (2018) covered only the first one to two years after implementation and therefore captured only short-run impacts. Third, I extend the heterogeneity analysis beyond gender and race to include age, education, and marital status, showing that the decline in job switching is concentrated among older workers nearing retirement who may place greater value on health insurance coverage.

The remainder of the paper is organized as follows. Section 2 reviews the related literature. Section 3 provides institutional background on the Affordable Care Act. Section 4 describes the data used in the analysis. Section 5 outlines the empirical strategy, including the Difference-in-Differences framework, the event study, and the staggered DID approach. Section 6 presents the main results and a series of robustness checks. Section 7 concludes.

2 Related Literature

Leading up to the passage of the ACA, advocates argued that expanding public coverage would reduce dependence on employer-sponsored insurance and alleviate job lock. However, early empirical evidence on the relationship between employment-based insurance and job mobility has been mixed. Some researchers find employer based insurabce leads to a decrease in job turnover rates, which is evidence of job lock (Gruber and Madrian 1994; Monheit and

Cooper 1994; Dey and Flinn 2005; Hamersma and Kim 2009); while others find no evidence of job lock (Kapur 1998; Gilleskie and Lutz 1999; Berger et al. 2004). Barkowski (2020) provides evidence that access to public health insurance can reduce both job lock and job push.

Alongside this, a different strand of literature examines the impact of the ACA Medicaid expansion on labor supply. Leung and Mas (2018) investigates the effect of Medicaid expansion on employment, hours of work, and wages for low-income childless adults in states that chose to expand Medicaid. They find no effect on wages, hours worked, or part-time employment, but do find an increase in full-time employment. Kaestner et al. (2017) explores the same question using a sample that includes both parents and non-parents. They implement a Difference-in-Differences approach and a Synthetic Control Method (SCM) and find that Medicaid expansion has a positive but small and statistically insignificant effect on labor supply. Aslim (2016) studies the impact of Medicaid expansion on childless adults' employment transitions using monthly CPS data. They estimate a difference-in-discontinuities model and find that the medicaid expansion increased part-time employment relative to full-time work but had no significant effect on overall labor force participation.

Research on state-specific Medicaid policies has examined their impact on employment and labor supply. Garthwaite et al. (2014) examines the rollback of Medicaid eligibility in Tennessee in 2005. Using a Difference-in-Differences design with Tennessee as the treated state and other Southern states as controls, they find mixed results. For low-educated, childless adults, the policy change led to a 25 percent increase in employment, while no effect was found for other education groups. Dague et al. (2017) studies the expansion of Medicaid to childless adults in Wisconsin in 2009. Using a quasi-experimental regression discontinuity design that exploited an enrollment cap preventing eligible individuals from enrolling after a certain date, they find that Medicaid enrollment was associated with a 2 to 18 percent decrease in employment. Baicker et al. (2013) examines Oregon's 2008 Medicaid expansion to childless adults and finds that gaining coverage was associated with a small, statistically insignificant decrease in employment and earnings.

The studies most closely related to this study are by Gooptu et al. (2016) and Callison and Sicilian (2018). Gooptu et al. (2016) uses CPS monthly data from January 2005 to March 2015 to evaluate the effect of Medicaid expansion on adults in expansion states, focusing on transitions from full-time to part-time work, transitions from employment to unemployment, and job switching. They define job switching as equal to one if an individual, employed in both years of their CPS panel, reported having just switched jobs in the second CPS interview but not in the first. In contrast, I define job switching on a month-to-month basis, as this captures short-term labor market adjustments that are missed in lower-frequency measures, reduces recall error, and allows for more precise detection of responses to policy changes. Using a Difference-in-Differences framework, they find no significant effect of Medicaid expansion on any of these outcomes. Callison and Sicilian (2018) also use CPS monthly data to study the effect of Medicaid expansion on labor force participation, employment, hours worked, and job mobility. They examine adults aged 18–64, a group that includes younger individuals covered by the ACA dependent coverage mandate and parents who may gain insurance through their children's eligibility. Their results show that Medicaid expansion has no effect on job mobility but increases labor force participation and employment.

3 Background

The Affordable Care Act (ACA) was enacted in 2010 with the aim of significantly expanding health insurance coverage across the United States. Before the ACA, public health insurance programs were limited to certain groups, which mainly included the disabled, low-income children, pregnant women, and some parents who met strict income limits. Most workers in the US rely on employer-sponsored insurance[2] though not all jobs offer such benefits.

[2]According to the Annual Social and Economic Supplement for 2010, 55.3% of Americans or 169 million people have employer-sponsored health insurance. In 2023, the American Community Survey (ACS), reports

Individuals who do not qualify for public insurance and lack access to employer-based plans often face a high cost of purchasing insurance in the private market.

To address these gaps, the ACA implemented a broad set of reforms aimed at making health insurance more accessible and affordable for a wider range of people. These reforms included individual and employer mandates to encourage participation in the insurance system, the creation of online health insurance exchanges where consumers could compare and purchase plans, the introduction of income-based subsidies to lower premiums and out-of-pocket costs, and a major expansion of Medicaid eligibility to cover more low-income adults. Together, these changes sought to reduce the number of uninsured Americans and reshape the health insurance landscape.

A key feature of the ACA was the expansion of Medicaid to cover low-income adults with incomes up to 138% of the Federal Poverty Line (FPL). In addition, the program provided subsidies to help individuals earning between 138% and 400% of the FPL afford coverage in the private market. Under the original design of the ACA, all states and the District of Columbia were expected to expand Medicaid, with the federal government covering the full cost of expansion in 2014 and gradually reducing its share to 90% after five years. However, the Supreme Court ruled that this requirement was unconstitutional, allowing states to decide whether to adopt the expansion. As a result, twenty-six states and the District of Columbia chose to expand Medicaid in 2014, followed by three more in 2015 and another two in 2016. Table 1 presents the list of states and their Medicaid expansion dates as of 2019.

Before the ACA, state Medicaid programs only provided coverage to low-income parents of dependent children and offered no coverage to the rest of the non-disabled adult population. States that wanted to extend coverage to groups not required under federal law had to apply for a "demonstration" waiver, which required states to prove that the expansion would not increase federal Medicaid spending relative to what it would have been without

this figure to be 53.7 %

the waiver. In 2013, as reported by the Kaiser Family Foundation, eight states[3] and the District of Columbia offered insurance coverage to childless adults under specific income thresholds. The ACA expansion represented a major policy shift by formally including this group in the eligibility criteria.

According to the American Community Survey (ACS), Medicaid enrollment rose by approximately three million individuals between 2013 and 2014. A substantial share of Medicaid recipients were childless adults: in 2013, about 45% of respondents reported having no children in their household, and this share increased to 54% in 2014 and 55% in 2015. I use residence in a Medicaid expansion state as a natural experiment in this study.

4 Data

The primary data used in this study are the Current Population Survey (CPS) Basic Monthly data, obtained from IPUMS. State-level monthly unemployment rates are sourced from the Bureau of Labor Statistics (BLS). The analysis covers the period from January 2008 through December 2019. The CPS Basic Monthly data are nationally representative and contain detailed information on household demographics, labor market outcomes, and family income. Approximately 60,000 households are interviewed each month using a rotating panel design: households are surveyed for four consecutive months, then exit the sample for eight months, and return for a final four-month interview period. In addition, I use data from the American Community Survey (ACS) to document Medicaid take-up among non-parent adults in expansion and non-expansion states. I present these descriptive trends in the results section to illustrate the coverage effects of the policy.

The main sample includes eighteen states that expanded Medicaid in 2014 and seventeen non-expansion states. Eight states and DC in this group had comprehensive programs that covered childless adults and closely resembled the ACA. These eight states are omitted from

[3] These states include: Arizona, Colorado, Connecticut, Delaware, Hawaii, Minnesota, New York and Vermont

the main analysis but are included in a robustness specification. Kaestner et al. (2017) added these states in the control group in their analysis, but I dropped them to avoid the confounding effects of Medicaid expansion. The states that expanded Medicaid after 2014 are incorporated in another robustness specification using a staggered adoption framework with the Wooldridge (2021) estimator, which I describe in a later section. Table 2 displays the expansion and non-expansion states in the main sample, the early and later treated states.

The sample is restricted to individuals ages 25 to 64. This age range excludes those under 25, who may be covered by the ACA's dependent coverage mandate enacted in 2010, and those 65 and older, who are generally eligible for Medicare. Military personnel are excluded from the analysis because they are covered by TRICARE, a military health insurance program. The CPS asks respondents in months two through four and six through eight of the rotation pattern whether they are still working for the same employer and job they reported as their main job in the previous month. Because this question is not asked in the first and fifth interview months, I exclude individuals in those months from the analysis. The sample consists of individuals who were employed in the previous month at the time of the interview.

Since the group of interest in this study is adults without dependent children, I restrict the sample to individuals who do not have any dependent children residing in their household. Adults with dependent children were more likely to qualify for Medicaid prior to the ACA, and including them could introduce bias if their monthly labor market transitions are systematically influenced by unobservable factors related to their pre-existing eligibility. The CPS Basic Monthly data report annual family income in $5,000 intervals. Following Leung and Mas (2018), we approximate family income using the upper bound of each income category. Individuals are classified as below the poverty line if the upper bound of their household income category is less than 138% of the Federal Poverty Line (FPL) for their family size in the interview year, based on the official poverty thresholds. As anticipated, this approach yields a conservative estimate of poverty. Callison and Sicilian (2018) use the midpoint of each income category. We also report results using this measure as a robustness check.

The dependent variable in this analysis is a monthly employer-switching indicator. It is coded as 1 if an individual reports working for a different employer than in the previous month, and 0 otherwise. This variable is constructed using the short panel structure of the CPS, which allows individuals to be tracked over time. We focus on individuals employed in consecutive months; this outcome captures the monthly probability of an individual changing employers. Thus, rather than measuring aggregate transitions, this approach allows for an individual-level analysis of short-run job mobility.

Our main estimation sample consists of 95,889 person-month observations, corresponding to 38,101 unique individuals. Table 3 reports weighted summary statistics for individuals in treated and control states, along with the differences in their means. Women represent about 44% of the population in both groups. The share of white individuals is 59% in treated states compared to 56% in control states, though this difference is not statistically significant. About 40% of individuals in treated states are between ages 26 and 39, compared to 35% in control states. Roughly 38% of respondents in both groups report having completed high school as their highest level of education.

To provide a clearer picture of pre-treatment comparability, Table 4 presents the same summary statistics restricted to the period before Medicaid expansion in 2014. I find that, with the exception of the share of Black individuals, the proportion of those aged 55–64, and the average unemployment rate, the treated and control states are similar across most observable characteristics. The similarity in these key groups supports the parallel trends assumption underlying the Difference-in-Differences framework. Nevertheless, I complement this with an event study analysis to further assess whether the assumption holds.

Table 4 shows that, in the pre-treatment period, the mean of the job-switch variable for the treated group is higher than the overall mean, indicating greater switching before the policy change. To illustrate this, Figure 1 plots the mean job-switch rates over the study period. After 2016, the mean for the treated group declines slightly, while the control group remains relatively constant. The pre-treatment trends also suggest parallel trajectories, as the treated and control groups closely follow the same trend prior to the policy

5 Identification

To estimate the causal effect of the ACA Medicaid expansion on worker mobility, I compare changes in mobility between low-income workers in states that expanded Medicaid and those that did not, before and after the expansion. Specifically, I estimate the following Difference-in-Differences (DiD) specification:

$$Y_{ist} = \beta_0 + \beta_1 ME_{it} + \beta_2 X_{ist} + \beta_3 UR_{st} + \gamma_s + \rho_t + \varepsilon_{ist} \tag{1}$$

where $Y_{ist} = 1$ if individual i, residing in state s at time t, is no longer employed by the same employer or job as in the previous month, and 0 otherwise. The vector γ_s captures state fixed effects, while ρ_t accounts for month-year fixed effects. UR_{st} denotes the state-level unemployment rate, and X_{ist} is a vector of individual-level covariates. The variable of interest is ME_{it}, an indicator equal to one if individual i resides in a state that had expanded Medicaid coverage to childless adults by time t, and zero otherwise. β_1 captures the average effect of Medicaid expansion on the probability that an individual switches employers from one month to the next, relative to individuals in non-expansion states.

The vector of individual characteristics includes gender; age groups (25-39, 40-54, and 55-64); race/ethnicity (non-Hispanic White, non-Hispanic Black, Hispanic, and other); educational attainment (less than high school, high school graduate, some college, and college graduate); and marital status. All regressions use final-sample weights provided by the CPS, and robust standard errors are clustered at the state level. Equation (1) represents the main specification. To examine heterogeneous effects across subpopulations, we interact the group indicator with the Medicaid expansion variable ME_{it} in the following way:

$$Y_{ist} = \beta_0 + \beta_1 ME_{it} + \beta_2 Z_i + \beta_3 (ME_{it} \times Z_i) + \beta_4 X_{ist} + \beta_5 UR_{st} + \gamma_s + \rho_t + \varepsilon_{ist} \qquad (2)$$

where Z_i denotes the subgroup indicator (e.g., Female, Age(55-64), Married, High-school or less, and White). The coefficient β_1 captures the effect of Medicaid expansion on the baseline group (those with $Z_i = 0$), while β_3 measures the differential effect for subgroup Z_i relative to the baseline. The other variables are as previously defined.

A key identifying assumption of the difference-in-differences (DID) approach is that job mobility in expansion and non-expansion states would have followed similar trends in the absence of Medicaid expansion. Any deviation from this parallel trend assumption after the policy change is attributed to the effect of the expansion. I am also interested in estimating the dynamic effects of Medicaid expansion on job mobility over time. As stated earlier, it is possible that the effect of the policy does not occur immediately but accumulates or dissipates gradually in the months following implementation. To test this assumption and examine the dynamics of the treatment effect over time, I estimate an event study specification of the form:

$$Y_{ist} = \sum_{\tau=-t_{\min}}^{t_{\max}} \beta_\tau \cdot D_{st}^{\tau} + \delta X_{ist} + \gamma_s + \rho_t + \varepsilon_{ist} \qquad (3)$$

In this specification, D_{st}^{τ} is an indicator equal to 1 if individual i in state s is observed τ months relative to the state's Medicaid expansion month (e.g., $\tau = -1$ denotes one month before expansion, $\tau = 0$ denotes the year of expansion, and so on). In my main sample, I set $\tau = -1$ to be December 2013. The coefficients β_τ trace the dynamic path of the treatment effect over time.

All event-time indicators are defined relative to the omitted category $\tau = -1$, which serves as the baseline. Coefficients for $\tau < 0$ test for pre-trend differences between expansion and non-expansion states, while coefficients for $\tau \geq 0$ capture the dynamic effects of Medicaid expansion on job switching behavior. The model includes individual controls X_{ist}, state fixed effects γ_s, and month-year fixed effects ρ_t, with standard errors clustered at the state level.

5.1 Staggered DID

Five states expanded Medicaid after 2014, as shown in Table 1. Earlier studies in the Medicaid expansion literature either excluded these states (Gooptu et al. 2016), added them to the control group (Moriya et al. 2016), or used the traditional two-way fixed effects (TWFE) framework (Kaestner et al. 2017; Leung and Mas 2018). As Goodman-Bacon (2021) shows, TWFE can yield biased estimates when treatment timing is staggered.[4]. To incorporate later-treated states without this bias, I estimate treatment effects using Wooldridge (2021) difference-in-differences (DiD) estimator, which is robust to staggered adoption. Let Y_{ist} denote an indicator equal to 1 if individual i in state s at month t switched employers relative to the previous month, and 0 otherwise. For each expansion cohort g (the year of Medicaid adoption), I estimate:

$$Y_{ist} = \sum_s \alpha_s \mathbf{1}\{S = s\} + \sum_t \lambda_t \mathbf{1}\{T = t\} + \sum_g \beta_g D^g_{st} + X_{ist}\gamma + \varepsilon_{ist}, \tag{4}$$

where $\mathbf{1}\{S = s\}$ and $\mathbf{1}\{T = t\}$ are indicator variables for state and month-year fixed effects, respectively. The term X_{ist} is the same vector of controls defined in Equation (1), and $D^g_{st} = \mathbf{1}\{s \in g,\, t \geq G\}$ equals 1 if state s belongs to cohort g (adoption year G) and the month t is after adoption. The coefficient β_g captures the average treatment effect for cohort g by comparing treated states in that cohort to never-treated states.

The overall average treatment effect on the treated (ATT) is then:

$$ATT = \sum_g w_g \, \beta_g, \tag{5}$$

where $w_g = \dfrac{N_g}{\sum_g N_g}$ are weights proportional to the number of treated observations N_g

[4]Traditional TWFE estimators can yield biased estimates when treatment timing is staggered because comparisons between early- and late-treated units can produce negative weighting

in cohort g.

6 Results

6.1 Main results

Before looking at job mobility, it is useful to confirm that the Medicaid expansion actually affected coverage in my sample of childless adults. Figure 2 shows the share of Medicaid recipients in expansion and non-expansion states using ACS data. Before 2014, some states covered a limited number of childless adults through Section 1115 waivers or other state-funded programs, which explains the small positive level of coverage observed prior to expansion. Starting in 2014, coverage for childless adults rose sharply in the expansion states while remaining mostly flat in the control states. Between 2013 and 2019, Medicaid participation increased by 27.7 percent in expansion states compared to 7.2 percent in non-expansion states. This clear divergence illustrates the direct impact of the Affordable Care Act's Medicaid expansion.

Table 5 presents the basic regression results where we assess the effect of Medicaid expansion on the probability of switching jobs among Medicaid-eligible, childless adults. Column (1) reports the baseline regression without any controls, and Column (2) adds a rich set of demographic controls. All specifications cluster standard errors at the state level.

The coefficient in Column (1) indicates that Medicaid expansion reduced job switching by 0.48 percentage points, and this effect is statistically significant at the 5% level. In Column (2), which includes a richer set of controls, the estimated effect is –0.005 and remains statistically significant at the 5% level. This implies that Medicaid expansion lowered the monthly probability of job switching among Medicaid-eligible, childless adults by about 0.5 percentage points, representing an 18.3% decline relative to the pre-treatment mean job-switching rate of 2.7%.

The negative effect in our results is consistent with a decline in "job push" following

Medicaid expansion. Before expansion, uninsured workers are likely to switch into jobs offering ECHI to obtain coverage. By expanding Medicaid eligibility, the policy reduced this tendency. Medicaid-eligible individuals could now obtain insurance without needing to move into ECHI jobs, decreasing mobility driven by health insurance needs. In contrast, the mechanism of reduced "job lock" would predict an increase in job switching. Our estimates suggest that the job-push channel dominated in this population.

6.2 Heterogeneity results

Table 6 presents subgroup results. Medicaid expansion reduces job switching among men, whites, married individuals, and workers with education beyond high school, whereas the corresponding estimates for women, non-whites, and less-educated workers are smaller and statistically indistinguishable from zero. These patterns point to heterogeneous policy impacts, although formal tests of differences across subgroups do not yield statistically significant results.

A plausible reason for finding no effect among women is that they are more likely to obtain health insurance through other household arrangements, such as spousal coverage. As a result, Medicaid expansion would not alter their behavior in the job push context. By contrast, men are more often the primary insurance holders in households and would otherwise need to seek employer-sponsored coverage. Medicaid expansion reduces this pressure to search for a job that provides health insurance

The most notable heterogeneity arises by age. Near-retirees (ages 55–64) exhibit a sizable decline in monthly job switching following Medicaid expansion, whereas younger workers show no significant response. This differential effect indicates that older workers reduce mobility once Medicaid lowers dependence on employer-sponsored coverage, while younger workers are more likely to continue switching jobs after Medicaid eligibility is expanded. This can be explained by the fact that younger workers continue to seek jobs that better match their skills, which may result in higher wages in the long run. In essence, younger workers' job mobility is less tied to the availability of health insurance.

6.3 Event studies

I present the event study results of the main analysis in Figure 3. The estimates show that the confidence intervals in the pre-treatment period are centered around zero, supporting the validity of the parallel trends assumption. After the policy takes effect, the dynamics indicate that Medicaid expansion does not significantly reduce job switching in the short run, though coefficients are largely negative. The decline in mobility only becomes apparent three years after the expansion. At that point, the magnitude of the effect reaches about half a percentage point, consistent with the DiD estimates. This pattern is consistent with Gooptu et al. (2016) and Callison and Sicilian (2018), who examined the first fifteen months and the first two years after implementation, respectively, and found no significant effects. My findings suggest that the impact of Medicaid expansion on job switching may take longer to emerge.

Figure 4 shows the event study results by gender and age group. The pre-treatment trends look flat, which means the parallel trends assumption holds for these groups. After the expansion, women and prime-age workers show no change in job switching. In contrast, men and older workers show a delayed drop in job switching, which is similar to the timing of the effect in the main sample.

Figure 5 presents the event studies by race, education, and marital status. Married individuals, whites, and workers with education beyond high school show a decline in job switching, while unmarried individuals, non-whites, and less-educated workers show little or no change. Importantly, for the groups that do respond, the decline emerges only gradually in the years following expansion rather than immediately. This suggests that the policy's impact on mobility reflects an adjustment process as households learn about eligibility, navigate enrollment, and adapt their employment decisions over time.

6.4 Alternative Samples

Table 7 reports the estimates separately for parents and non-parents within the Medicaid-eligible population in the original sample states. The goal is to test whether parents were also affected by the expansion. The results show no effect of Medicaid expansion on parents' job-switching behavior, likely because many parents already had access to public health insurance prior to the ACA expansion. Consistent with this, I find no significant effects when considering the entire Medicaid-eligible population (parents and non-parents combined). This result agrees with Gooptu et al. (2016), which also used a sample consisting of parents and non-parents.

I estimate the effect of the ACA Medicaid expansion on eligible individuals in states that had similar pre-existing programs but nonetheless expanded in 2014. The idea is to see whether the broader eligibility rules and the stronger federal financing of the ACA created additional labor market effects even in environments where some coverage was already in place. Following the specification sequence from Table 5, I begin with a baseline regression without controls, then add demographic controls. Across these models, we find that the ACA Medicaid expansion reduced the probability of month-to-month job switching by about 0.8 percentage. This suggests that even in states with prior Medicaid-like programs, the ACA expansion generated meaningful changes, which likely reflects that the federal expansion expanded eligibility and provided a more credible, stable source of coverage compared to the limited pre-ACA programs.

To obtain a complete set of early adopters and avoid bias from selectively excluding states that had ACA-like programs before 2014, I combine the early expansion states with those that expanded in 2014. This approach ensures that all states where low-income childless adults gained Medicaid eligibility by 2014 are included in the treatment group, providing a fuller and more accurate picture of the policy's effect on job-switching. Table 9 reports results for this sample. The estimates are consistent with my main findings: Medicaid expansion reduces the monthly probability of job switching by about 0.6 percentage points.

In Table 10, I present the estimates for all states that expanded Medicaid between 2014 and 2016, using a staggered DiD framework. For comparison, I also report results from the conventional TWFE specification, and both approaches yield consistent findings, showing that Medicaid expansion is associated with a reduction in monthly job transitions of approximately 0.48 percentage points under the Wooldridge estimator and 0.46 percentage points under the TWFE model. Estimating the effect for all expansion states provides a more complete picture of the policy's overall labor market impact, rather than focusing only on early adopters.

6.5 Robustness Tests

I carry out a number of robustness tests, including sensitivity and falsification exercises. In the sensitivity test, I re-estimate the model using alternative federal poverty cutoffs instead of 138% in order to verify that our findings are not driven by the precise eligibility threshold chosen. For robustness, I also replicate the main analysis using the sample definition from Kaestner et al. (2017)[5], which provides an alternative treatment-control classification. These exercises help ensure that the results do not only hold to one particular specification or sample definition, but instead reflect a more general underlying relationship. For the falsification test, I restrict the sample to individuals ages 16–25, who were already covered by the ACA's dependent coverage mandate and therefore should not be affected by Medicaid expansion.

Table 11 presents the regression results using alternative FPL cutoffs. Restricting the sample to individuals below 150% of the FPL, I find that Medicaid expansion reduces monthly job transitions by 0.6 percentage points, a result very similar to the baseline sample. Using a broader cutoff of 200% FPL, the estimated effect is a 0.3 percentage point reduction in job switching. The pattern suggests that the policy's impact is strongest among the lowest income groups, who are most likely to gain coverage from Medicaid, while the effect diminishes and loses statistical significance as income rises and individuals become less

[5] Our baseline follows Leung and Mas (2018), who exclude states with either comprehensive pre-ACA Medicaid expansions or limited prior expansions for non-parent Medicaid-eligible adults.

directly affected by the expansion.

Table 12 reports the results of the falsification test. Using the Kaestner et al. (2017) sample, the estimates indicate that Medicaid expansion reduced job switching by 0.5 percentage points, which is similar to the effect found in the main sample. In contrast, the results show no effect of the ACA Medicaid expansion on monthly job transitions among individuals aged 16 to 25, consistent with the expectation that this group was already covered under the ACA dependent coverage mandate and therefore was not directly affected by the expansion.

7 Discussion

The findings in this paper contribute to the broader discussion of whether Medicaid expansion affected labor market mobility. While earlier studies such as Gooptu et al. (2016) and Callison and Sicilian (2018) examined different, broader populations and found no significant effects, the results here show that meaningful impacts emerged only among childless adults and became evident after a three-year adjustment period. Specifically, we find that Medicaid expansion reduced the probability of job switching among Medicaid-eligible, childless adults by roughly 0.5 percentage points, corresponding to an 18.5% decline relative to the pre-treatment mean. This pattern is more consistent with a reduction in job push than with a reduction in job lock

Another reason our work differs from Gooptu et al. (2016) is our use of a monthly transition definition, which captures more frequent switches and can detect short-term changes that their two-year recall window may miss. The results also differ from Callison and Sicilian (2018), who analyze a sample that includes parents and young adults ages 16–25. Because these groups had access to public health insurance prior to Medicaid expansion, their labor market behavior is not expected to respond strongly to the policy. To examine whether the delayed effects I document for childless adults extend to this population, I replicate the analysis of Callison and Sicilian (2018) using data extended through 2019. Table 13 reports the results, which show that Medicaid expansion had no significant effect on this population, confirming that their labor market behavior was largely unresponsive to the policy.

An important feature of my results is their timing. The effect of Medicaid expansion does not appear immediately but emerges only after a lag of roughly three years. This delayed response highlights that labor market adjustments to major health policy reforms can take time, as workers and households gradually respond to new incentives.

My study is not without limitations. Because the CPS does not directly observe employer-sponsored insurance, the estimates are intent-to-treat effects and cannot fully distinguish between job lock and job push mechanisms. Intent-to-treat estimates provide an appropriate framework for assessing the causal impact of Medicaid expansion on eligible populations, and this is the approach we adopt in this paper

8 Conclusion

This research examines the effect of the Affordable Care Act's (ACA) Medicaid expansion, implemented in 2014, on the probability of individual job-switching. The analysis focuses on non-elderly, non-disabled, childless adults with household income below 138 percent of the federal poverty line. The sample includes individuals living in states that adopted the expansion in 2014 and did not have pre-existing ACA-like policies. The result of the estimation shows that Medicaid expansion reduced the likelihood of job switching by about 0.5 percentage points.

The findings lend support to the job-push hypothesis, which suggests that uninsured workers are more likely to seek jobs that offer health insurance, even if those jobs pay lower wages. Medicaid expansion changes this incentive by reducing the need for low-income workers to switch jobs in search of coverage. The results show that childless, low-income workers are less likely to change jobs when a valuable benefit missing from their current employment, such as health insurance, becomes available through public policy.

The heterogeneity analysis shows that the effect of Medicaid expansion is concentrated among older workers. Among all the subgroups examined, the effect of Medicaid expansion differs significantly between older workers (ages 55–64) and prime-age workers (ages 26–54). A plausible explanation for this pattern is that younger workers may place less value on health insurance when switching jobs, as they are more likely to change jobs for better wages or career opportunities. In contrast, older workers tend to care more about their health and the security of health insurance, so when coverage is publicly provided, they have less incentive to switch jobs.

These results contrast with earlier studies in the literature, which generally found no effect of Medicaid expansion on job mobility. One explanation is that prior work focused on the early post-expansion years, when it may have been too soon for Medicaid eligibility changes to influence labor market behavior, as shown in Gooptu et al. (2016) Another is that other research analyzed populations outside the main target of the expansion, including groups less likely to gain coverage, as reported by Callison and Sicilian (2018).

Overall, the findings highlight that the effects of major health policy reforms on job mobility can take several years to materialize, a consideration that is important for evaluating future public policy.

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Graphs

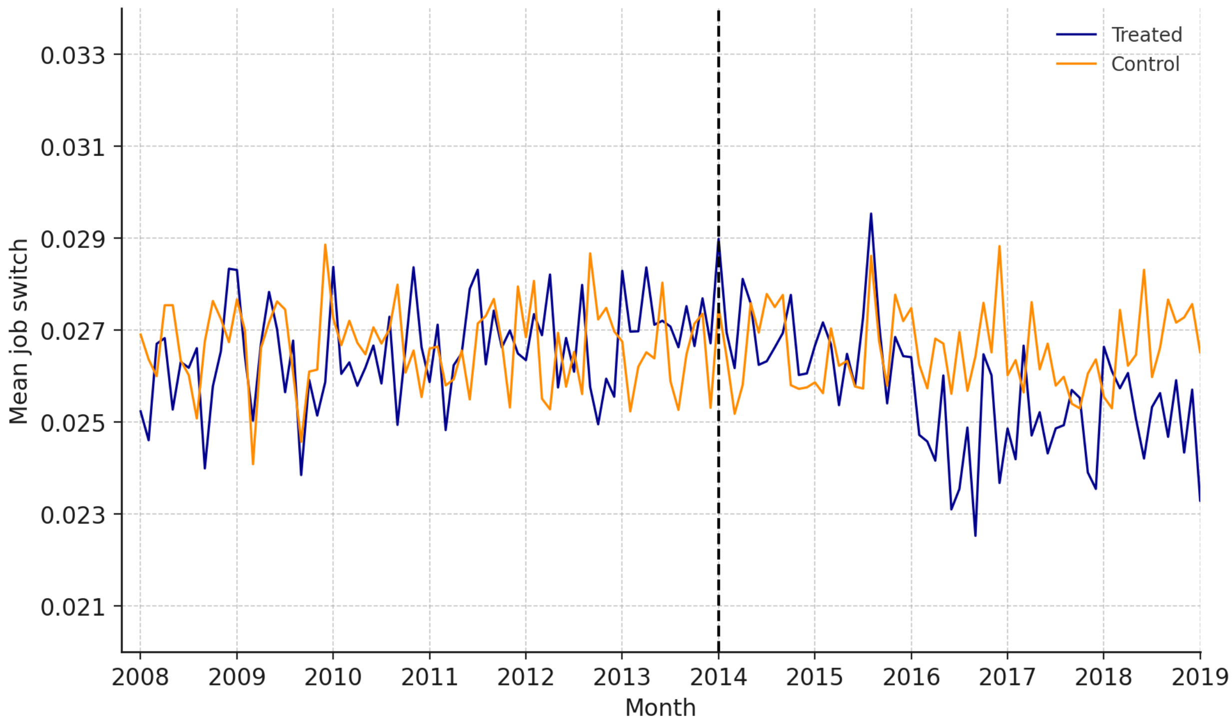


Figure 1: Means of monthly job-switch for Childless Adults: Expansion Vs Non-Expansion States

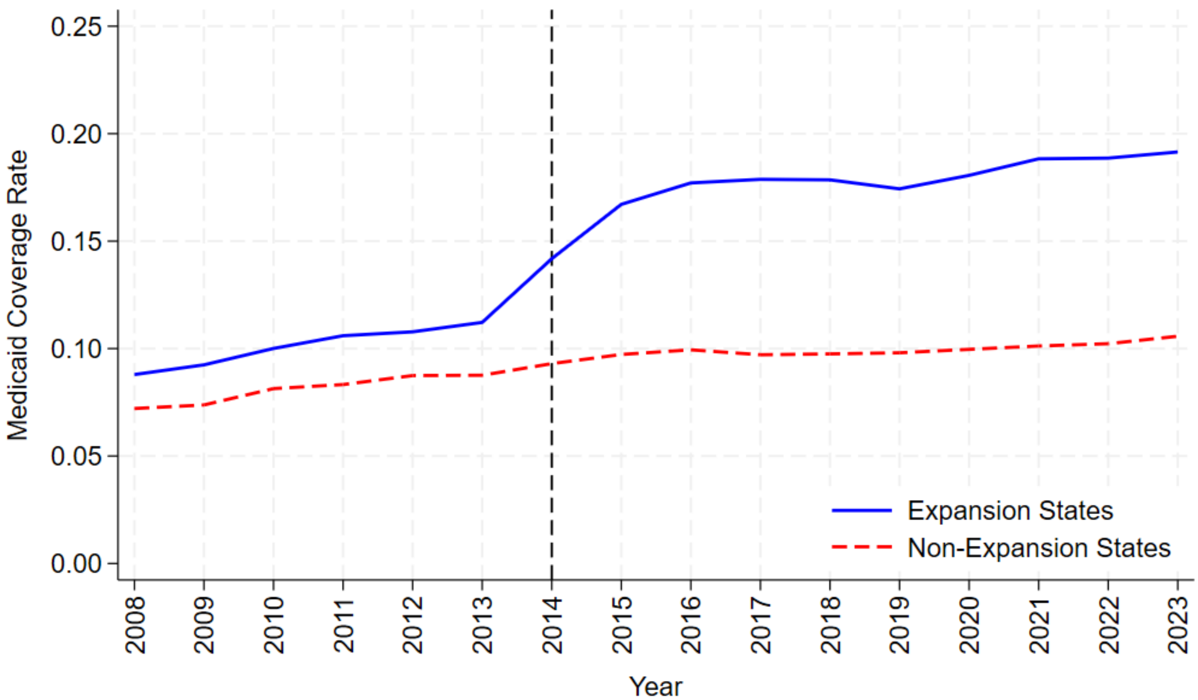


Figure 2: Medicaid Coverage for Childless Adults: Expansion Vs Non-Expansion States

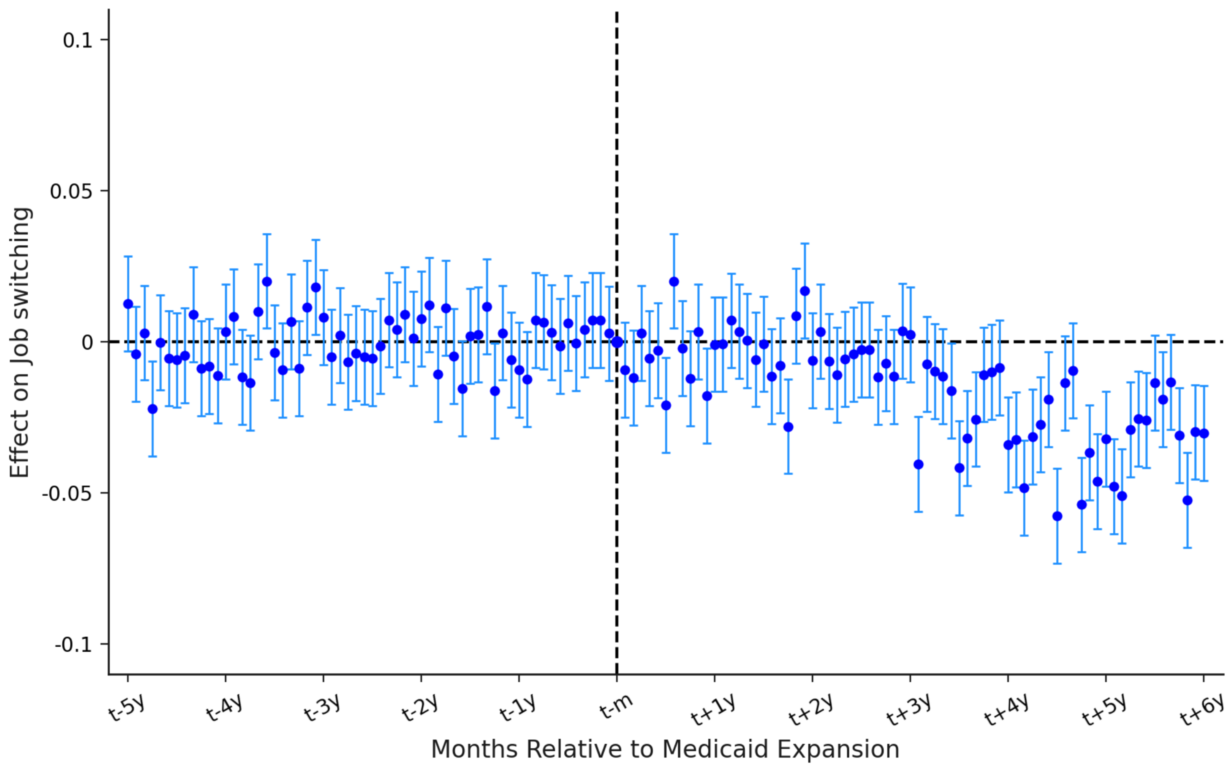
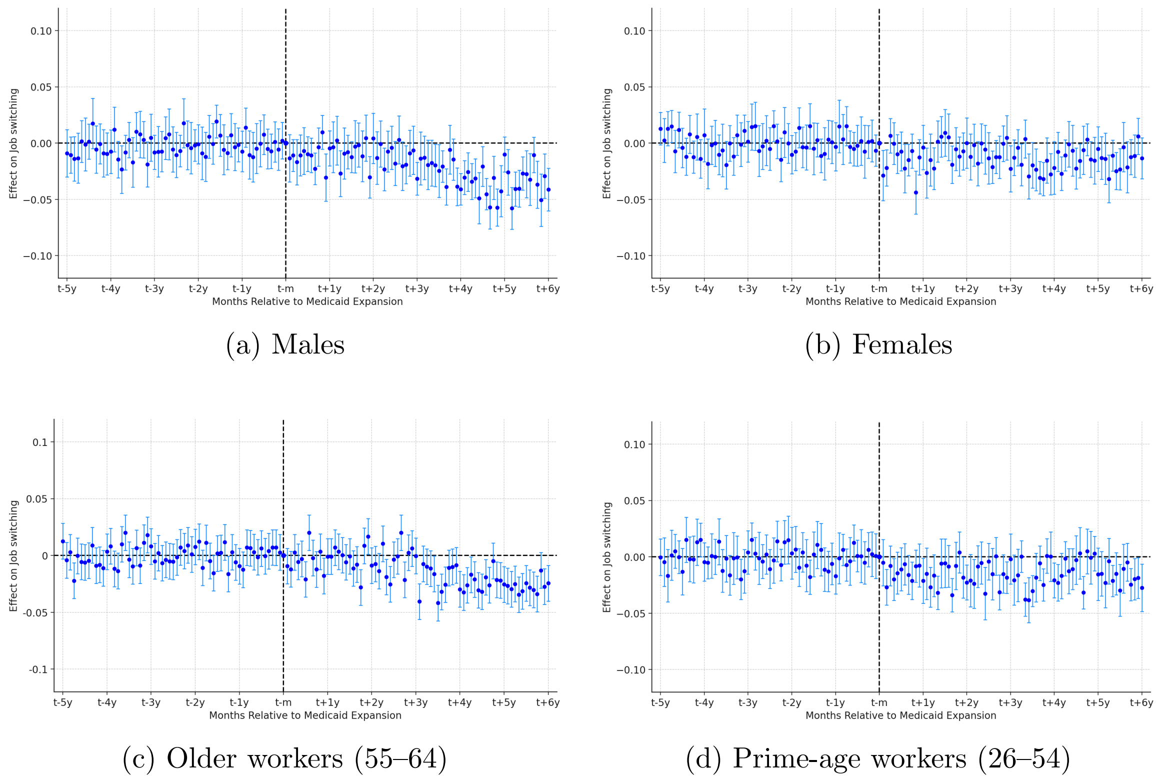


Figure 3: Event study estimates of Medicaid Expansion on job mobility of non-parents



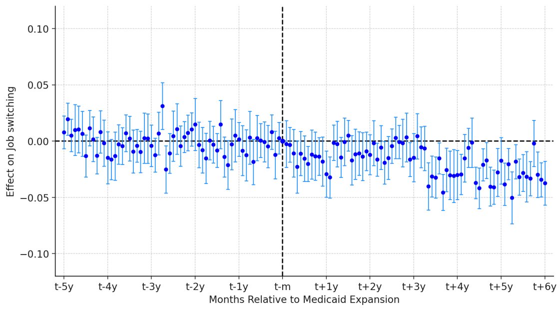
(a) Males

(b) Females

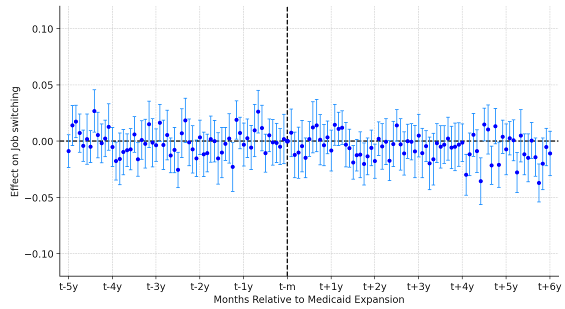
(c) Older workers (55–64)

(d) Prime-age workers (26–54)

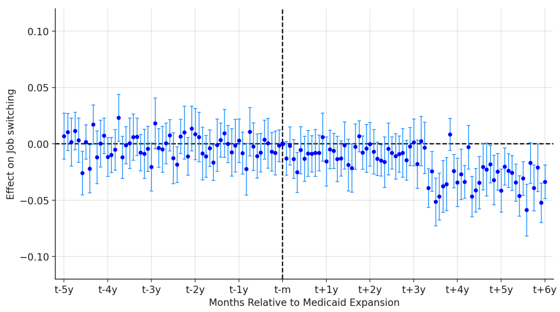
Figure 4: Event study estimates for Medicaid expansion: heterogeneity by gender and age group.



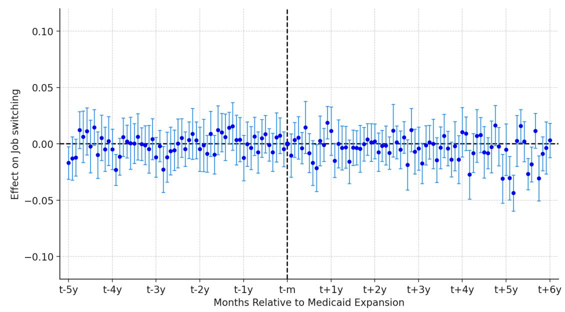
(a) Whites



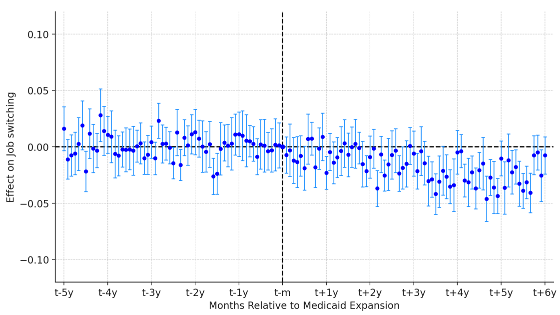
(b) Non-whites



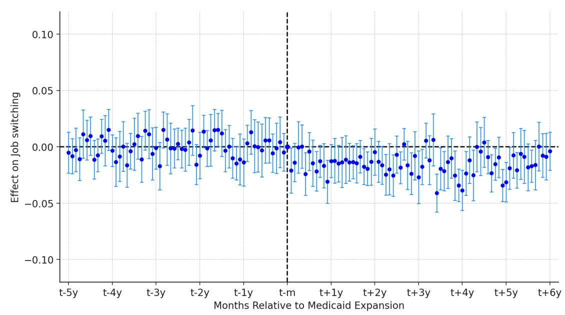
(c) Above High school



(d) High school or less

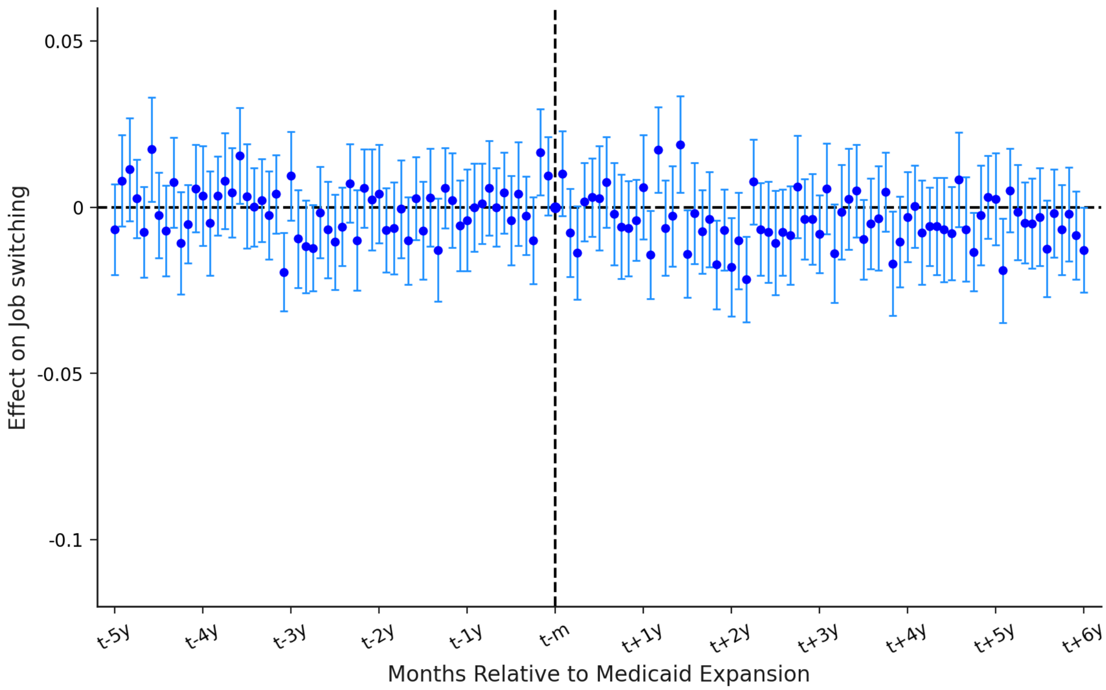


(e) Married

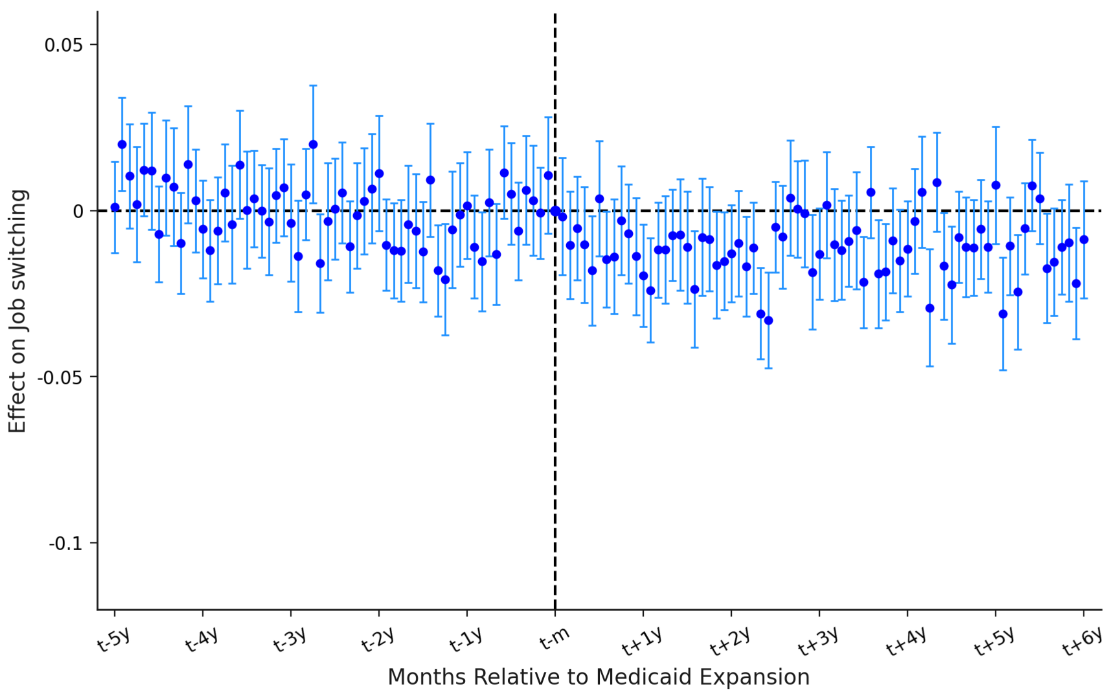


(f) Unmarried

Figure 5: Event study estimates for Medicaid expansion: heterogeneity by race, education, and marital status.



(a) Parents



(b) Medicaid-eligible

Figure 6: Event study estimates for Medicaid expansion on parents and the entire eligible population

Tables

Table 1: Classification of States by Medicaid Expansion Status

Expanded by 2014	Expanded After 2014	Non-Expansion States
Arizona (1/1/2014)	Pennsylvania (1/1/2015)	Alabama
Arkansas (1/1/2014)	Indiana (2/1/2015)	Florida
California (1/1/2014)	Alaska (9/1/2015)	Georgia
Colorado (1/1/2014)	Montana (1/1/2016)	Idaho
Connecticut (1/1/2014)	Louisiana (7/1/2016)	Kansas
Delaware (1/1/2014)		Maine
District of Columbia (1/1/2014)		Mississippi
Hawaii (1/1/2014)		Missouri
Illinois (1/1/2014)		Nebraska
Iowa (1/1/2014)		North Carolina
Kentucky (1/1/2014)		Oklahoma
Maryland (1/1/2014)		South Carolina
Massachusetts (1/1/2014)		South Dakota
Michigan (4/1/2014)		Tennessee
Minnesota (1/1/2014)		Texas
Nevada (1/1/2014)		Utah
New Hampshire (8/15/2014)		Virginia
New Jersey (1/1/2014)		Wisconsin*
New Mexico (1/1/2014)		Wyoming
New York (1/1/2014)		
North Dakota (1/1/2014)		
Ohio (1/1/2014)		
Oregon (1/1/2014)		
Rhode Island (1/1/2014)		
Vermont (1/1/2014)		
Washington (1/1/2014)		
West Virginia (1/1/2014)		

* Wisconsin did not adopt the ACA Medicaid expansion, but implemented a state-funded policy covering adults up to 100% FPL while leaving those between 100–138% FPL to access marketplace subsidies. Following prior studies, we classify Wisconsin as a non-expansion state.

Table 2: State Treatment Status by Medicaid Expansion Timing

Treated States (18)	**Control States (19)**	**Early Expansion (8)**	**Late Expansion States (5)**
(Expanded Jan 2014)	*(Never Expanded)*	*(ACA-like Coverage)*	*(Expanded after 2014)*
Arkansas	Alabama	Arizona	Alaska
California	Florida	Colorado	Indiana
Illinois	Georgia	Connecticut	Louisiana
Iowa	Idaho	Delaware	Montana
Kentucky	Kansas	District of Columbia	Pennsylvania
Maryland	Maine	Hawaii	
Massachusetts	Mississippi	Minnesota	
Michigan	Missouri	New York	
Nevada	Nebraska	Vermont	
New Hampshire	North Carolina		
New Jersey	Oklahoma		
New Mexico	South Carolina		
North Dakota	South Dakota		
Ohio	Tennessee		
Oregon	Texas		
Rhode Island	Utah		
Washington	Virginia		
West Virginia	Wisconsin		
	Wyoming		

Notes: States are grouped by timing of Medicaid expansion under the ACA.

Table 3: Summary Statistics of the main sample

	Treated	Control	Difference
Job switch	0.0250	0.0280	-0.003
	(0.1641)	(0.1650)	
Female	0.4396	0.4484	-0.0088
	(0.4963)	(0.4973)	
White	0.5904	0.5617	0.0287
	(0.4918)	(0.4962)	
Black	0.1004	0.2067	-0.1063**
	(0.3006)	(0.4050)	
Other race	0.2140	0.1805	0.0335*
	(0.4101)	(0.3846)	
Hispanic	0.0952	0.0511	0.0441
	(0.2934)	(0.2202)	
Age 26–39	0.4093	0.3552	0.0541**
	(0.4917)	(0.4786)	
Age 40–54	0.3401	0.3713	-0.0312
	(0.4737)	(0.4832)	
Age 55–64	0.2507	0.2736	-0.0229
	(0.4334)	(0.4458)	
No HS	0.1746	0.1956	-0.0210
	(0.3796)	(0.3967)	
High school	0.3765	0.3991	-0.0226
	(0.4845)	(0.4897)	
Some college	0.2578	0.2579	-0.0001
	(0.4374)	(0.4375)	
College	0.1911	0.1474	0.0437*
	(0.3932)	(0.3545)	
Married	0.2938	0.3164	-0.0226
	(0.4555)	(0.4651)	
Unemp. Rate	7.0013	6.1872	0.8141**
	(2.7109)	(2.2779)	
Observations	45,184	50,705	

Means reported; standard deviations in parentheses. CPS sampling weights are applied. Differences in means between treated and control groups are tested using two-sided t-tests. Reported p-values indicate the statistical significance of these differences.

Table 4: Summary Statistics of the main sample (Pre-Expansion, 2008–2013)

	Treated	Control	Difference
Job switch	0.0272	0.0261	0.0011
	(0.1596)	(0.1563)	
Female	0.4212	0.4299	-0.0087
	(0.4938)	(0.4951)	
White	0.5410	0.4908	0.0502
	(0.4983)	(0.4999)	
Black	0.1143	0.2331	-0.1188**
	(0.3181)	(0.4228)	
Hispanic	0.0990	0.0454	0.0536
	(0.2986)	(0.2082)	
Other race	0.2458	0.2307	0.0151
	(0.4306)	(0.4213)	
Age 26–39	0.4264	0.3665	0.0599
	(0.4945)	(0.4818)	
Age 40–54	0.3478	0.3888	-0.0410
	(0.4763)	(0.4875)	
Age 55–64	0.2258	0.2447	-0.0189*
	(0.4181)	(0.4299)	
No HS	0.2046	0.2319	-0.0273
	(0.4034)	(0.4220)	
High school	0.3655	0.3953	-0.0299
	(0.4816)	(0.4889)	
Some college	0.2482	0.2389	0.0093
	(0.4320)	(0.4264)	
College	0.1817	0.1339	0.0479
	(0.3856)	(0.3405)	
Married	0.2917	0.3238	-0.0321
	(0.4546)	(0.4679)	
Unemp. Rate	9.1828	8.0559	1.1268**
	(2.1707)	(1.9233)	
Observations	24,251	27,165	

Means reported; standard deviations in parentheses. CPS sampling weights are applied. Differences in means between treated and control groups are tested using two-sided t-tests. Reported p-values indicate the statistical significance of these differences.

Table 5: Effect of Medicaid Expansion on Job Switching

	(1)	(2)
Medicaid expansion	-0.0048**	-0.0051**
	(0.0022)	(0.0022)
Female		-0.0028**
		(0.0010)
Age 26–39		0.0058***
		(0.0014)
Age 40–54		0.0023
		(0.0014)
White		0.0028
		(0.0019)
Black		0.0020
		(0.0026)
Hispanic		0.0049**
		(0.0021)
High School		-0.0049***
		(0.0014)
Some College		-0.0021
		(0.0017)
College		-0.0015
		(0.0019)
Married		-0.0063***
		(0.0010)
Unemployment Rate		-0.0007
		(0.0009)
State FE	✓	✓
Month-Year FE	✓	✓
Observations	95,889	95,889
R-squared	0.0024	0.0033

Notes: Robust standard errors clustered by state in parentheses. All regressions include state and month-year fixed effects. Sample restricted to nonelderly, nondisabled, childless adults below 138% FPL.* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Effect of Medicaid Expansion on Job Switching by Subgroups

	(1) Gender	(2) Race	(3) Education	(4) Marital Status	(5) Age Group
Male	-0.0054**				
	(0.0026)				
Female	-0.0045				
	(0.0031)				
Female - Male	0.0009				
	(0.0024)				
White		-0.0056**			
		(0.002)			
Non-White		-0.0043			
		(0.0032)			
Non-White - White		-0.0012			
		(0.0032)			
Above High School			-0.0069**		
			(0.0023)		
High School or Less			-0.0034		
			(0.0029)		
$>$HS - $\leq$HS			0.0034		
			(0.003)		
Unmarried				-0.0045	
				(0.003)	
Married				-0.0065**	
				(0.0025)	
Unmarried - Married				0.002	
				(0.0028)	
Age(55 - 64)					-0.0081**
					(0.0025)
Age(25 - 54)					-0.0039
					(0.0024)
Age(25 - 54) - Age(55 - 64)					-0.0042*
					(0.0022)
Observations	95,889	95,889	95,889	95,889	95,889
State FE	✓	✓	✓	✓	✓
Month-Year FE	✓	✓	✓	✓	✓

Notes: Each column reports estimates from separate regressions of monthly job switching on the Medicaid expansion indicator interacted with subgroup dummies for gender, race, education, marital status, and age. The reported coefficients represent intent-to-treat effects for each subgroup. Rows labeled as differences (e.g., Female – Male) report the coefficient on the interaction term, testing whether effects differ significantly across groups. All regressions include state and month-year fixed effects. Standard errors are clustered by state.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 7: Effect of Medicaid Expansion on Job Switching: Parents and All Adults

	(1)	(2)	(3)	(4)
Medicaid	0.0022	0.0018	0.0001	-0.0003
	(0.0017)	(0.0016)	(0.0020)	(0.0015)
Female		-0.0038***		-0.0039***
		(0.0009)		(0.0006)
Age 26–39		0.0078***		0.0063***
		(0.0011)		(0.0008)
Age 40–54		0.0028***		0.0019**
		(0.0010)		(0.0009)
White		0.0020		0.0030
		(0.0021)		(0.0018)
Black		-0.0005		0.0008
		(0.0024)		(0.0019)
Asian		0.0036		0.0044**
		(0.0022)		(0.0018)
Other Race		-0.0006		0.0044
		(0.0047)		(0.0036)
HS Graduate		-0.0020		-0.0027**
		(0.0014)		(0.0011)
Some College		-0.0004		-0.0008
		(0.0012)		(0.0009)
College		0.0010		-0.0006
		(0.0012)		(0.0010)
Married		-0.0036***		-0.0048***
		(0.0008)		(0.0005)
Unemployment Rate		-0.0011*		-0.0010*
		(0.0005)		(0.0005)
State FE	✓	✓	✓	✓
Month-Year FE	✓	✓	✓	✓
Observations	217,779	217,779	313,668	313,668

Notes: Columns (1)–(2) report estimates for the sample of parents, while Columns (3)–(4) present results for all nonelderly adults. Columns (2) and (4) include demographic and labor market controls. All regressions include state and month-year fixed effects. Robust standard errors are clustered by state. *Significance levels:* * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 8: Regression results for only early treated states

	(1)	(2)
Medicaid expansion	-0.0078***	-0.0075***
	(0.0024)	(0.0024)
Female		-0.0021
		(0.0015)
Age 26–39		0.0054***
		(0.0016)
Age 40–54		0.0004
		(0.0015)
Age 55–64		[omitted]
Race: White		0.0054*
		(0.0029)
Race: Black		0.0059
		(0.0039)
Race: Hispanic		0.0097**
		(0.0037)
Other Race		0.0080
		(0.0063)
High School		-0.0055***
		(0.0018)
Some College		-0.0029*
		(0.0017)
College		-0.0025
		(0.0023)
Married		-0.0052***
		(0.0012)
Unemployment Rate		-0.0006
		(0.0015)
State FE	✓	✓
Month-Year FE	✓	✓
Observations	69,429	69,429
R-squared	0.0032	0.0041

Robust standard errors clustered by state in parentheses.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

This table shows the estimates for the states that had a similar policy to the ACA and also expanded Medicaid in 2014. They include: Arizona, Colorado, Connecticut, Delaware, Hawaii, Minnesota, New York, Vermont, and Washington D.C.

Table 9: Effect of Medicaid Expansion on Job Switching: All 2014 Expansion States

	(1)	(2)
Medicaid expansion	-0.0057***	-0.0058***
	(0.0020)	(0.0020)
Female		-0.0022**
		(0.0011)
Age 26–39		0.0057***
		(0.0013)
Age 40–54		0.0027**
		(0.0013)
Age 55–64		[omitted]
Race: White		0.0061***
		(0.0017)
Race: Black		0.0063**
		(0.0024)
Race: Hispanic		0.0095***
		(0.0022)
Other Race		0.0102*
		(0.0051)
High School		-0.0048***
		(0.0014)
Some College		-0.0016
		(0.0015)
College		-0.0006
		(0.0017)
Married		-0.0055***
		(0.0009)
Unemployment Rate		-0.0009
		(0.0008)
State FE	✓	✓
Month-Year FE	✓	✓
Observations	114,613	114,613
R-squared	0.0021	0.0030

Robust standard errors clustered by state in parentheses.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

This table shows the estimates for all states that expanded Medicaid in 2014. This includes states that had state health insurance programs covering non-parents before the ACA as well as those that extended eligibility to non-parents in 2014.

Table 10: Effect of Medicaid Expansion on Job Switching

	Staggered DiD	TWFE
Medicaid Expansion (Overall ATT)	-0.0048**	-0.0046**
	(0.002)	(0.002)
Group-Specific ATTs (Staggered DiD only)		
2014 Expansion States	-0.0052***	
	(0.0019)	
2015 Expansion States	0.0023	
	(0.0024)	
2016 Expansion States	-0.0058**	
	(0.0023)	
Observations	126,189	126,189
R^2	0.0035	0.0029

Notes: Table reports the effect of Medicaid expansion on job switching for all the states that extended eligibility to childless adults between 2014 to 2016. Staggered DiD results come from `jwdid`, while TWFE is a two-way fixed effects model. Standard errors (in parentheses) are clustered at the state level. Significance levels: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 11: Effect of Medicaid Expansion on Job Switching by FPL Cutoff

	(1)	(2)	(3)
	≤138% FPL	≤150% FPL	≤200% FPL
Medicaid expansion	-0.005**	-0.006**	-0.003*
	(0.002)	(0.002)	(0.002)
Female	-0.0028**	-0.0027**	-0.0026***
	(0.0011)	(0.0010)	(0.0008)
Age 26–39	0.0058***	0.0059***	0.0067***
	(0.0014)	(0.0013)	(0.0012)
Age 40–54	0.0024*	0.0026*	0.0035***
	(0.0014)	(0.0013)	(0.0011)
High school	-0.0049***	-0.0045***	-0.0040***
	(0.0014)	(0.0013)	(0.0011)
Some college	-0.0023	-0.0020	-0.0011
	(0.0017)	(0.0016)	(0.0013)
College	-0.0015	-0.0013	-0.0001
	(0.0019)	(0.0018)	(0.0014)
Married	-0.0062***	-0.0060***	-0.0049***
	(0.001)	(0.001)	(0.001)
Unemployment rate	-0.001	-0.001	-0.001
	(0.001)	(0.001)	(0.001)
State FE	✓	✓	✓
Month-Year FE	✓	✓	✓
Observations	95,889	100,994	172,497
R-squared	0.0034	0.0029	0.0024

Robust standard errors clustered by state in parentheses.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Table 12: Effect of Medicaid Expansion on Job Switching: Kaestner et al. Sample and Falsification Test (Ages 16–25)

	(1) Kaestner et al.	(2) Ages 16–25
Medicaid expansion	-0.005**	0.002
	(0.002)	(0.002)
Female	-0.003**	-0.001
	(0.001)	(0.002)
Age 26–39	0.007***	-
	(0.001)	
Age 40–54	0.003**	-
	(0.001)	
White	0.005**	0.008**
	(0.002)	(0.004)
Black	0.004	0.002
	(0.003)	(0.004)
Asian	0.007***	0.003
	(0.002)	(0.003)
High School	-0.005***	0.005
	(0.001)	(0.003)
Some College	-0.002	0.001
	(0.002)	(0.002)
College	-0.002	0.005
	(0.002)	(0.004)
Married	-0.005***	0.003
	(0.001)	(0.004)
Unemployment Rate	-0.001	-0.002**
	(0.001)	(0.001)
State FE	✓	✓
Month-Year FE	✓	✓
Observations	102,738	58,121
R-squared	0.003	0.007

Notes: Column (1) reports estimates for the Kaestner et al. (2017) analytic sample, restricted to nonelderly adults ages 26–64. Column (2) presents a falsification test for younger adults ages 16–25 who were largely ineligible for Medicaid under the ACA. All regressions include state and month-year fixed effects, and standard errors are clustered by state. *Significance levels:* * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 13: Callison and Sicilian (2018) Estimates

	Full Sample	138% FPL
Panel A: 2008–2019		
Medicaid Expansion	0.001	0.003
	(0.001)	(0.002)
Observations	4,870,610	544,907
Panel B: 2008–2016		
Medicaid Expansion	0.001	0.002
	(0.001)	(0.002)
Observations	3,746,512	425,061
State Fixed Effects	✓	✓
Month–Year Fixed Effects	✓	✓

Notes: This table replicates the results of Callison and Sicilian (2018). The lower panel (2008–2016) reproduces their original estimates, while the upper panel extends their sample period through 2019. All regressions include state and month–year fixed effects, and robust standard errors clustered by state are shown in parentheses. *Significance levels:* *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.