MEDICAID-ING COVERAGE VOLATILITY: EVIDENCE FROM THE 2019 VIRGINIA MEDICAID EXPANSION*

Bradley Heim,[†] Ithai Lurie,[‡] Elena Patel,[§] Shanthi Ramnath[¶]

ABSTRACT

In this paper, we leverage novel administrative tax data to study the effects of expanded access to Medicaid following a job loss. We present detailed summary statistics describing individual coverage dynamics of more than 1.6 million individuals who separate from an employer plan, including covariates associated with post-separation uninsurance spells. We exploit the late adoption of expanded Medicaid in Virginia to provide estimates of the Medicaid expansion on health insurance coverage dynamics. Using differences-in-differences, we estimate that access to Medicaid increases the likelihood of finding coverage by 17% and reduces the duration of uninsurance by 12%.

Keywords : Health Insurance Dynamics, Unemployment, Medicaid Expansion

JEL Classification : J65, I13, I18, I38

*This research was conducted while the authors were employees at the U.S. Department of the Treasury, the Federal Reserve Bank of Chicago, or the Federal Reserve System. The findings, interpretations, and conclusions expressed in this paper are entirely those of the authors and do not necessarily reflect the views or the official positions of the U.S. Department of the Treasury, the Council of Economic Advisors, the Federal Reserve Bank of Chicago, or the Federal Reserve System. All taxpayer data used in this research were kept in a secured Treasury or IRS data repository, and all results have been reviewed to ensure that no confidential information is disclosed. We thank conference participants and discussants at the Midwestern Economics Association, the American Society of Health Economists, the National Tax Association, the Texas A&M Microeconomics Empirical Workshop, and the Michigan Tax Invitational Conference. In addition, we thank seminar participants at the University of Chicago, the Federal Reserve Bank of Chicago, Brigham Young University, the University of Utah, the University of Massachusetts at Lowell, the University of Exeter, the University College Dublin, and the Congressional Budget Office. Finally, we thank Mike Geruso and Amy Cyr for valuable feedback throughout the development of this work.

[†]O'Neill School of Public and Environmental Affairs, Indiana University, heimb@indiana.edu

[‡]Office of Tax Analysis, U.S. Treasury, 1500 Pennsylvania Ave NW, Washington DC, 20220, ithai.lurie@treasury.gov

[§]David Eccles School of Business, University of Utah, 1731 South Campus Center Drive Suite 3400; Salt Lake City, UT 84112, elena.patel@eccles.utah.edu

[¶]Federal Reserve Bank of Chicago, 230 South LaSalle Street Chicago, IL 60604, Federal Reserve Bank of Chicago, sramnath@frbchi.org

1. Introduction

Did the Affordable Care Act's (ACA) historic expansion of Medicaid enlarge the program's focus from public assistance to additionally include providing stop-gap health coverage to a broader population? In this paper, we study how the Medicaid expansion affects coverage volatility for those who involuntarily lose their private employer health insurance due to job loss. Coverage volatility compromises individual health and elevates the risk of financial setbacks (see, for example, Cutler and Gelber, 2009; Guevara et al., 2014; Gai and Jones, 2020), which can be especially salient following a sudden job loss. While access to Medicaid mitigates such coverage volatility, many households were ineligible prior to the ACA's expansion of Medicaid benefits. Yet, the risk of coverage volatility is one facing the general population as most Americans access health insurance through employment (Lurie and Pearce, 2021), and as a result, face coverage uncertainty due to cyclical labor markets. The ACA's Medicaid expansion created scope to considerably ease this volatility because the new eligibility guidelines extended benefits to many who experience job loss; but, whether the Medicaid expansion benefited this broader population remains an open question.

For our analysis, we make use of novel administrative tax data that reports *monthly* information on the source of health insurance coverage at the individual level for the U.S. population. In 2016 alone we observe roughly 11 million individuals who separate from their employer plan, 1.6 million of whom are unemployed. We document several stylized facts about insurance transitions among the unemployed population, expanding our understanding of coverage dynamics that had been previously unobservable on a large-scale basis. We then estimate the effect of expanded Medicaid access on coverage dynamics by leveraging the quasi-natural variation provided by the 2019 Virginia Medicaid expansion. Overall, our work uses millions of observations to highlight the expanded role of Medicaid as a crucial component to a broad social safety net.

We focus on the population of policyholders aged 18 to 62 who lose employer coverage and simultaneously claim unemployment benefits; this latter restriction focuses our analysis on those more likely to have lost both their employment and insurance coverage unexpectedly.[1] In 2016, our data include more than 1.6 million policyholders, for whom we observe longitudinal monthly coverage data spanning the twenty-four months following their coverage loss. We then link to individual tax return data to capture information about the policyholder, including employment status, earnings, age, marital status, gender, and geographic location.

We use these data to characterize coverage dynamics among those experiencing unexpected job and coverage loss. First, the average duration of uninsurance is 4.9 months. Second, 6 percent of those dropping coverage fail to regain coverage within two years, implying the existence of some longer run persistence in uninsurance. Third, Medicaid serves as the first source of new coverage for nearly one-quarter of those who become unemployed. Fourth, those living in states with expanded Medicaid access are 37 percent more likely to regain coverage, all else equal.

Having established these baseline statistics, we next leverage quasi-natural policy variation to estimate the effect of the ACA's Medicaid expansion on both the likelihood and duration of uninsurance. We study the 2019 Virginia Medicaid Expansion using a difference-in-differences model that compares post-separation coverage dynamics of (1) those separating from an employer policy in Virginia before and after the Medicaid expansion to (2) those who separated concurrently in states without expanded Medicaid.[2] This setting allows for

[1]As we explain later, we characterize policyholders as unemployed if they receive unemployment income, reported on Form 1099-G, in the year of or in the year after they separate from an employer plan.

[2]During our sample period, two states expanded Medicaid: Virginia and Louisiana. Due to data limitations impacting Louisiana, we focus our analysis on the Medicaid expansion in Virginia.

an estimation the effects of increasing Medicaid access to the unemployed while simultaneously controlling for other time-invariant factors that could also influence coverage dynamics. We find that the Virginia Medicaid expansion increased the likelihood of regaining coverage within the following year by 17% and reduced the duration of uninsurance by 12%. Individuals are 1.5 times more likely to transition to Medicaid as a first source of coverage in Virginia due to the expansion. Further, the effects are concentrated among low-wage workers who were 33% more likely to regain coverage in the year of their job loss.

Finally, we explore heterogeneity by gender and marital status as these demographic characteristics affect the likely impact of a Medicaid expansion. In particular, married individuals who are eligible for spousal coverage have an additional means through which to obtain insurance, and therefore, may be less responsive to increased Medicaid access. Women, however, have higher utilization of health care (Bertakis et al., 2000) and, therefore, may be more impacted by increasing Medicaid access. Consistent with this, we find that both margins are important for how increased access to Medicaid impacts coverage dynamics. Women are 21% more likely to find coverage by the end of the year compared to men, who are 13% more likely to find coverage. Similarly, unmarried individuals are 25% more likely to find coverage compared to the 5% estimate for married individuals, which is not statistically significant.

Overall, our research uses high-quality administrative data covering the U.S. population to highlight how the ACA's Medicaid expansion provides stop-gap coverage to a broader population. Insurance instability has long been a concern among policymakers; plan transitions and gaps in coverage are costly from both a health and a monetary standpoint (Roberts and Pollack, 2016; Gai and Jones, 2020; Kressin et al., 2020). While a deep empirical literature studies the effects of Medicaid — including the topic of insurance instability — this body of work has relied on small-scale survey data and is primarily focused on how this public assistance program benefits the chronically low-income population (see, for example, Swartz, Marcotte and

McBride, 1993; Swartz and McBride, 1990; Fairlie and London, 2008; Schaller and Stevens, 2015; Schaller and Zerpa, 2019; East and Simon, 2022). We show that Medicaid also provides benefits to those who lose their job and, as a result, access to their employer plan. In this way, the ACA widened the Medicaid safety net through its ability to help stabilize coverage for the majority of Americans who typically rely on their employers for health insurance and perhaps also by playing a role in broader employment decisions where the unemployed are able be more selective in their job pursuits.

2. Background

2.1. Sources of Health Insurance Coverage in the US

The two largest sources for public health insurance in the United States are Medicare, which serves as a single-payer insurance system for adults above 65, and Medicaid, which provides no-cost insurance to eligible, low-income adults and their children. Medicare is managed by the federal government, whereas Medicaid is managed in collaboration between federal and state governments. This co-administration of Medicaid leads to varying eligibility standards across states, which are determined by factors such as monthly income, assets, and household composition.

Given Medicaid's eligibility criteria, which is based on monthly income that can be volatile, and the frequent eligibility determinations conducted by state Medicaid offices, low-income individuals are observed to "churn" in and out of eligibility over time. For example, the Kaiser Foundation estimates that during 2018 — nearly a decade after the Affordable Care Act (ACA) was signed into law — roughly 10% of Medicaid enrollees were disenrolled from Medicaid but subsequently re-enrolled within one year (Corallo et al., 2021). In some cases, individuals may have what is referred to as "latent eligibility" for Medicaid. This term captures those individuals who meet the eligibility criteria for Medicaid but who are not enrolled or aware of their eligibility at the time they receive medical services. In this case, if they subsequently enroll in Medicaid, or if a social worker or case worker completes enrollment on their behalf, their coverage can be retroactively extended to cover medical expenses incurred prior to application during periods of latent eligibility.

The majority of Americans are covered by private insurance, which is most commonly accessed through employers who sponsor tax-subsidized group insurance for their employees.[3] Access to such health insurance policies are generally offered as part of a compensation package, and employers may choose to pay a portion (or all) of the plan's monthly premium. Individuals who are not offered access to employer health plans and are not eligible for public insurance plans can purchase coverage in what is known as the non-group market.

The ACA notably introduced financial incentives to allow states to expand access to Medicaid by increasing the eligibility threshold to 138% of the Federal Poverty Level for all adults; 25 states introduced a Medicaid expansion concurrent with the other provisions of the ACA that came online in 2014. The ACA introduced a host of new regulations, including that large employers must offer affordable insurance coverage to their employees or else face a penalty. In addition, the ACA expanded access to non-group private insurance by creating subsidized marketplaces where individuals without access to an employer plan and ineligible for Medicaid could purchase insurance.

Similar to Medicaid, coverage instability is a risk for individuals with private insurance. However, instability in private coverage, which includes both intensive margin (who provides coverage and the amount of coverage) and extensive margin changes (whether someone has

[3]We define employer health plans as coverage obtained through one's own or one's spouse's employer, including multi-employer plan and Small Business Health Options Program (SHOP) plans.

any coverage) changes, likely results from a number of different channels. Most notably for this population, labor market dynamics, including voluntary or involuntary job changes, can force an employee to separate from their private employer policy. Indeed, prior to the ACA's enactment, job loss was associated with a nearly 20 percentage point increase in the subsequent likelihood of uninsurance. (Schaller and Stevens, 2015; Schaller and Zerpa, 2019; East and Simon, 2022). Alternatively, if an employee's hours are reduced (voluntarily or not), then they may no longer qualify for health coverage despite remaining employed. Employees may also choose to change their plan type or insurer during open enrollment or may switch to dependent coverage through a spouse or partner; each of these adjustments will result in coverage instability on the intensive margin.

2.2. Related Literature: Insurance Instability and Coverage Dynamics

A number of past studies have explored the extent of insurance instability in the U.S, especially prior to the enactment of the ACA. These studies use survey data and focus on the characteristics and duration of uninsurance spells (Swartz, Marcotte and McBride, 1993; Swartz and McBride, 1990; Fairlie and London, 2008; Einav and Finkelstein, 2023; Schaller and Stevens, 2015; Schaller and Zerpa, 2019; East and Simon, 2022). Early work from Swartz and McBride (1990), using data from the 1984 Survey of Income and Program Participation (SIPP), finds that roughly half of all uninsured spells end within four months . Cutler and Gelber (2009), using SIPP data covering the periods 1983–1986 and 2001–2004, show that the likelihood of losing any coverage grew from 19.8% to 21.4% from the 1980s to the early 2000s . They additionally show that shorter periods of uninsurance were associated with transitions to public insurance in the early 2000s.

More recent work studies changes in coverage dynamics that arose after the enactment of the the ACA. Using data from the Medical Expenditure Panel Survey (MEPS), Graves and Nikpay (2017) show that transitions made by the uninsured to private and public coverage increased after the health care expansion. Vistnes and Cohen (2018) uses the Household Component of the MEPS (MEPS-HC) and shows that uninsurance spell duration declined in 2014–15, after the implementation of the ACA, relative to 2012–2013. Agarwal and Sommers (2020) use MEPS data to measure the effects of the ACA on those experiencing job loss before and after the policy change. Finally, Gai and Jones (2020) use the MEPS to describe early changes in insurance instability across different types of coverage with the implementation of the ACA.

An important caveat of these studies lies in the data limitations of survey data, which generally preclude a causal analysis of the effect of the ACA. For instance, the MEPS survey is designed to be *nationally* representative rather than representative at the state level. Consequently, these data are ill-suited for conducting state-level analyses, where significant variation in Medicaid access exists. Moreover, the sample size of any single MEPS panel, consisting of roughly 13,000 families and 30,000 individuals, is too small to permit an analysis of sudden policy separations resulting from job loss — a meaningful trigger for insurance loss (Fairlie and London, 2008; Schaller and Stevens, 2015; Schaller and Zerpa, 2019; East and Simon, 2022).

The SIPP provides an alternative source of longitudinal health insurance coverage data. However, this survey, while similar in sample sizes to the MEPS, suffers from a well-known "seam bias." This bias occurs when changes in coverage within the reference period are underreported, and too many transitions occur between interview rounds. It is particularly serious when studying duration data, which is the focus of this paper (Ham, Li and Shore-Sheppard, 2009).

To address these limitations, we contribute to this literature by leveraging novel administrative tax data. This newly available data includes coverage reporting and state variation, allowing us to provide causal estimates of the effect of Medicaid expansion for those who lose health coverage due to unemployment.

3. Identifying Employer Plan Separations Using US Tax Data

The Affordable Care Act imposed new regulations requiring insurers and employers to report individual health insurance coverage to the IRS. This information is documented on Forms 1095-A, 1095-B, and 1095-C. Specifically, Form 1095-A captures monthly coverage acquired through insurance policies purchased on the ACA-created state and federal marketplaces, Form 1095-B captures monthly coverage acquired through government programs like Medicare and Medicaid, and both Forms 1095-B and 1095-C capture monthly coverage acquired through private policies, including employer-provided health insurance.[4] Each form reports which individuals, both policy holders and dependent beneficiaries, are associated with a particular policy in a given tax year. We exploit the high-frequency and longitudinal nature of these data to provide a comprehensive analysis of private health coverage dynamics within the U.S.

3.1. Data Construction

Although reporting requirements began in tax year 2014, transition rules for the first year of the Affordable Care Act offered reporting relief. As such, we begin by identifying individuals

[4]We define employer plans based the presence of Form 1095-C, which is used by employers who qualify as an Applicable Large Employer (ALE), or based on the following 1095-B, line 8 codes: code A (Small Business Health Options Program); code B (Employer-sponsored coverage); or code E (multi-employer plans). See Lurie and Pearce (2021) for a more detailed description of these tax forms.

who separate from an employer plan in 2016 to allow one full year of observations prior to separation, and we construct an individual-level panel dataset describing monthly coverage information surrounding their plan separation.We require that individuals were covered by the same policy for at least twelve months prior to focus on well-attached policyholders. Conditional on twelve months of continuous enrollment, we then define an individual as having separated from their employer plan in month m if they are covered in month m according to a single Form 1095-B or Form 1095-C, but not covered in the next month, $m+1$.[5] Next, we combine information from all three 1095 Forms across multiple tax years to create a panel of monthly coverage for these individuals; this panel contains coverage information the twelve months prior to the employer plan separation through twenty-four months after. We classify those who are not identified on any Form 1095 in month m as uncovered.[6]

For all of our analyses, we condition the sample to policyholders between ages 18 and 62 at the time of separation. This ensures that policyholders who experience a separation do not subsequently transition to Medicare during their twenty-four month post-separation period.[7] Finally, we incorporate into our analysis information describing an individual's geographic location, unemployment compensation, and wages using additional tax data. Geographic location is determined based on the address information reported on the Form 1095s. Unem-

[5]Each form reports individual coverage for a single tax year, from January–December. Using these forms, we identify individuals who separate from their employer plan by comparing monthly coverage over the year. This sample identification comes with one limitation: we do not identify those who separate in December, because that would require a comparison of monthly coverage across two different tax years, which is outside of the scope of our data construction.

[6]When employment is terminated, covered individuals are permitted to continue health insurance coverage for a limited period of time on their original employer plan under COBRA (Consolidated Omnibus Budget Reconciliation Act). These former employees must opt-in within three months of termination to continue on the same plan and are responsible for *both* the employer and employee portion of their monthly premium. This coverage is known as COBRA coverage. COBRA coverage can continue for 18–36 months, depending upon the circumstances surrounding the termination of employment. Our data cannot not distinguish between months of COBRA coverage and months of coverage provided through employment because an individual's plan does not change under COBRA—the only change is to who pays for the premium. Our data will identify the month in which a previously covered employee moves to a new plan or becomes uncovered following a lapse in COBRA coverage.

[7]Information on age and gender are provided to the IRS by the Social Security Administration (SSA).

ployment compensation (UI benefits) is reported on Form 1099-G, and wage data is reported on Form W-2. Finally, we determine marital status, household size, and household income from information reported on Form 1040 in the year prior to a plan separation.

3.2. 2016 Summary Statistics

Column (1) of Table 1 reports summary statistics for the population of policyholders who separated from their employer plan in 2016 after having been covered by that plan for at least one year. These 10.6 million policyholders comprise roughly 8% of the 132 million people who were covered by an employer health plan in 2016 (Lurie and Pearce, 2021).

Overall, policyholders who separate from an employer plan are more likely to be male and unmarried. The majority are aged 26 to 44, when typical life-cycle changes create higher employment volatility. They earned an average of $62,093 in wages in the year prior to separation (2015); by comparison, the median *household* income that year was $56,516 (Proctor, Semega and Kollar, 2016). Finally, 15.3% claimed UI benefits.

Column (1) also describes households' average Modified Adjusted Gross Income (MAGI) as a share of the Federal Poverty Limit (FPL) in 2015.[8] This benchmark takes into account household composition and determines eligibility for both Medicaid and subsidized marketplace coverage. Individuals with a MAGI less than 138% of the FPL are eligible for Medicaid in expansion states, while individuals with a MAGI between 138% and 400% of the FPL are likely to be eligible for a Premium Tax Credit (PTC) that offset the cost of purchasing health insurance through the marketplace.[9] The distribution of MAGI as a share of FPL in the

[8]The Poverty Guidelines are issued each year by the Department of Health and Human Services (HHS) and are used for administrative purposes to determine eligibility for certain programs. In 2015, the poverty threshold for a single-person household was set to $11,770. A four-person household faced a poverty threshold of $24,250. Full details of the 2015 poverty guideline appeared in the *Federal Register* on January 22, 2015.

[9]Note that the PTC *amount* is based on the difference between a reference premium policy, known as the second lowest cost silver plan (SLCSP), and the maximum required contribution the household is needs to pay

year before a policy separation implies that roughly 10% of these policyholders would have qualified for Medicaid in an expansion state and 46.9% would have otherwise qualified for a Premium Tax Credit.

3.3. Survey Data on Coverage Dynamics

While these administrative data are a relatively new resource, a number of publicly available survey data have been used in the past to study health insurance coverage. In general, many cross-sectional surveys, including the American Community Survey, the Current Population Survey, and the National Health Interview Survey, provide a measure of coverage during a reference period— for example "ever covered in the past year." These point-in-time measures can mask considerable coverage instability that occurs throughout the year, making them less useful for studying high-frequency transitions (Gai and Jones, 2020).

To our knowledge, only two public use surveys, the Medical Expenditure Panel Survey (MEPS) and the Survey of Income and Program Participation (SIPP), contain *within year* longitudinal coverage information detailing the source of coverage (e.g., private or public). However, neither survey is well-suited to study the monthly coverage observed in the administrative data. In addition, these surveys lack detailed information at the plan level. Without such identifiers, within-type transitions including—for example, transitions from one employer plan to another—are unobservable. By comparison, administrative tax data are reported at the plan level, allowing for a more comprehensive analysis in general of coverage instability.[10]

In addition, the sample size of both the MEPS and SIPP surveys are several orders of magnitude smaller than the tax data, making it difficult to estimate statistics about population-

for premiums. Hence, some people in the 138% to 400% of FPL might not get PTC if their required contributions exceed the SLCSP.

[10]While the tax data contain most coverage transitions, certain within-insurer plan changes, such as moving from a high deductible plan to a Health Maintenance Organization offered by the same provider are unobserved.

level employer plan separation and post-separation coverage dynamics with precision. The 2016 MEPS data, for example, include roughly 7,000 employer plan policyholders aged 18–62, of which 700 experience a change in coverage that year. Similarly, the 2014 SIPP includes roughly 8,000 employer plan policyholders aged 18–62 in its fourth wave, with 750 of that group experiencing a coverage transition in 2016. As previously described, the SIPP also suffers from a well known "seam" bias, where respondents are asked to report monthly health insurance over a lengthy backward looking period and, as a result, tend to report more changes in coverage toward the beginning of each interview period (Gai and Jones, 2020). Administrative tax data, by comparison, identify nearly 11 million instances in which a policyholder separates from an employer plan each year and do not suffer from recall bias like seam bias. As a result, the tax permits a more in-depth analysis allowing for an exploration of results along different dimensions of heterogeneity.

Finally, both the MEPS and the SIPP are designed to be *nationally* representative, complicating analyses of state-based Medicaid expansions.[11] The tax data, which contain precise geographic identifiers, provide an opportunity to exploit these state-level experiments.

Section 4 starts by providing a more general descriptive of coverage dynamics for the population of those who likely experienced an exogenous loss in coverage, i.e. the unemployed. Then in Section 5, we will utilize state level variation to estimate the causal effects of increased Medicaid access on high frequency coverage dynamics based variation created by Virginia.

[11]We note that the 2014 SIPP includes a large enough sample to be representative at the state level for the four largest states (CA, NY, FL, and TX).

4. Uninsurance Dynamics After Job Loss

4.1. Policy Separations Due to Unemployment

The data described above includes *all* policyholders who separate from an employer plan, regardless of why the separation occurred. However, the decision to regain coverage can be endogenous to the reason for the policy separation itself. For example, policyholders may choose to drop coverage in coordination with changes to ex-ante expectations of their medical expenditures, which would in turn influence how quickly they regain coverage. While the data do not allow us to identify those who exogenously separate from their policy, we can identify a subsample who is more likely to have experienced an *unexpected* loss in coverage by leveraging unemployment benefits tax reporting.[12] Typically, employees are only eligible to claim unemployment benefits if they are terminated due to layoffs, changes in business conditions, or a business closure. We hypothesize that individuals who separated from their plan and also received unemployment income were less likely to have coordinated their policy separation.

We characterize our policyholders as unemployed if they received unemployment insurance income (UI) as observed through third party reporting (Form 1099-G) in the year of, or the year after, a policy separation. We include two years of UI reporting to allow for delays in filing unemployment claims that can push the receipt of benefits to the next calendar year. This is especially likely for those who lose their job closer to the end of the year.

[12]While mass layoffs or plant closures could presumably identify more plausibly exogenous job loss, these events are not ideally suited to our setting. First, mass layoffs are not observed in the tax data, and therefore must be measured with noise using a threshold change in Form W2 counts or some alternative metric. Second, employees affected by mass layoffs or plant closures are often able to negotiate a continuation of health insurance, making the change in coverage less exogenous than the job loss.

Unemployed policyholders are a subsample of the full population of policyholders who separate from their employer plan.[13] After imposing this unemployment restriction, our sample of separated policyholders in 2016 is roughly 1.63 million people. For the remainder of the paper, we will focus on this group to isolate plausibly exogenous changes in coverage status; note that Appendix B replicates all of our descriptive statistics using the full population for the interested reader. Column (2) of Table 1 presents summary statistics for the unemployed who lost their coverage. Compared to the full population (col 1), they are older, less likely to be female, less likely to be married, and earn less. After a policy separation, they are less likely to be covered the following month, have a longer duration of uninsurance, and are more likely to remain uncovered for the full 24-month post-separation period.

Columns (3) and (4) of Table 1 report these same statistics for the unemployed based on whether they live in expansion or non-expansion state, respectively, prior to losing coverage. In general, policyholders who separate from an employer plan in non-expansion and expansion states look similar with respect to their gender and age compositions. Policyholders are slightly more likely to to be married (i.e., a joint filer) in non-expansion states, though the average rates are relatively comparable. Those in expansion states on average earn similar wages to those in non-expansion ($56,784 compared to $53,453), but are slightly less likely to qualify for a Premium Tax Credit measured as a higher share who earns more than 400% of the Federal Poverty Limit (40.1% compared to 35.4%).

4.2. Exiting Uninsurance

Panel (a) of Figure 1 reports the share covered by an employer plan (darkest green), Medicaid (middle green), a marketplace plan (lightest green), and the share uncovered (grey) in each

[13] This sample will necessarily exclude unemployed policyholders who did not claim UI benefits. As such, should be considered as a subsample of the full unemployed population.

month over the 24 month post-separation window. This figure highlights three important facts. First, 34% of individuals who separate from an employer plan move to a new employer plan immediately; these transitions are often missed in survey data, which capture transitions across broad types of coverage (e.g., private, public) but do not report changes within coverage type. Second, Medicaid covers an additional 15% of policyholders immediately following dropped employer coverage. Third, there is a substantial reduction in uninsurance during the first year after a loss in coverage that slows considerably in the second year.

To complement the point-in-time measure of coverage, panel (a) of Table 5 shows the six-month transition matrix for the full sample of unemployed people who lose their coverage. Of those who start out as uninsured in a given month, 47% remain uninsured 6 months later, while 43% transition to an employer plan. Both employer plans and Medicaid exhibit considerable persistence, though persistence among employer plans is stronger (82% vs. 61%). Finally, among those who find coverage under a new employer plan, 12% end up uninsured six months later, highlighting the existence of quick turnovers.

To shed light on selection out of uninsurance, Table 2 describes policyholders based on the length of their uninsurance spell. Each column of Table 2 reports statistics for policyholders by the month in which new coverage is obtained. The last column reflects statistics for those who remain without new coverage for at least 24 months ("Never Covered"). These summary statistics highlight that as the duration of uninsurance increases, individuals are more likely to be male and unmarried. In addition, prior wages decline with spell duration, suggesting positive selection out of uninsurance. Perhaps most surprising, the majority of policyholders are employed in the year following separation regardless of their spell length (including the never covered), which implies a decoupling of coverage and employment. Finally, roughly 6% of unemployed policyholders remain without health coverage for at least 24 months despite the fact that nearly three-quarters (71.4%) of them find employment in the following year.[14] This suggests the existence of longer-run persistence in uninsurance following job loss.

Table 3 describes policyholders based on their first source of coverage. We show that most of those who turn to Medicaid as a first source of coverage are employed in the year after separation (83.7%), consistent with other data sources highlighting high employment rates of Medicaid recipients.[15] In fact, the lowest rate of post-separation employment is among those with marketplace coverage, though the majority (80.4%) of this group is employed.[16] Comparing across employer and Medicaid coverage, there appears to be sorting by gender, where women are more likely to use Medicaid as a first coverage source while men are more likely to move to employer coverage. By contrast, coverage through the exchange is roughly evenly split across genders. In addition, there is sorting across coverage types by marital status, where non-married policyholders are more likely to take-up Medicaid and married policyholders are more likely to take-up employer coverage. Finally, the duration of uninsurance for those who move to Medicaid is almost one month *shorter* than for those who move to either an employer plan or a marketplace plan.

4.3. Factors Associated with Exits from Uninsurance

We next formalize the descriptive statistics presented in Tables 2 and 3 using a regression framework to characterize the dynamic process of regaining insurance. Age and gender have implications for expected health costs, and income and marital status correlate with differences

[14] Note that many unemployed policyholders who transition out of uninsurance return to it at a later date, so that the point-in-time fraction of unemployed policyholders who are uninsured 24 months post-separation seen in 1 substantially exceeds the fraction who are uninsured for the entire 24 months.

[15] See, for example, https://www.kff.org/medicaid/issue-brief/understanding-the-intersection-of-medicaid-work-a-look-at-what-the-data-say/

[16] As a reminder, a policyholder can move to an employer plan without being employed if they move to a spouse or partner's employer plan.

in an individual's access to insurance. As such, we estimate the extent to which each of these factors is associated with when to obtain insurance coverage, all else equal. We include an indicator for living in a Medicaid expansion state prior to dropping coverage as the expansion dramatically increases eligibility for Medicaid (Frean, Gruber and Sommers, 2017).

We estimate the following OLS model:

$$Y_i = \mathbf{X}_i\beta + \phi_m + u_i,$$

Here, Y_i can be one of several outcomes: (1) an indicator for obtaining coverage one month after separation, (2) the duration of uninsurance, measured in months, or (3) the likelihood of moving to either an employer plan, a Medicaid policy, or a marketplace plan as a first-source of coverage, conditional on regaining coverage within the 24 month post-separation period.[17] $\mathbf{X}_i$ captures demographic characteristics of the policyholder i: gender, filing status, wages, age, and access to Medicaid expansion. Filing status, wages, and access to the Medicaid expansion are all measured in 2015, the year prior to separation. All specifications include fixed effects for the month in which the policy separation occurred, ϕ_m. In addition, Appendix B presents estimates using the full population of policyholders who separate from an employer plan.[18]

We compliment these estimates using a Cox Proportional Hazard model, which can more directly account for the unique features of duration data. In particular, the probability of regaining insurance is unlikely to remain constant over time, and our finite post-separation observation period (24 months) induces right-censoring in our data: we do not observe the duration of uninsurance for the 6% of individuals who are uninsured for at least 24 months.

[17] As is common when studying dynamic processes, duration is right-censored data; persons who do not regain coverage after 24 months are coded as having a duration of uninsurance of 24 months. As described shortly, we explicitly account for this data anomaly using a survival model.

[18] Because the full population sample is not restricted to the unemployed, we include a dummy variable in the regression that identifies unemployment insurance receipt in 2016 or 2017.

The Cox Proportional Hazard model accounts for these data features and measures the impact of observable characteristics on the hazard ratio, i.e., the probability of finding coverage within an additional unit of time conditional on not having found coverage to that point. Appendix C provides additional background information on the Cox Proportional Hazard Model.

Columns (1) and (2) of Table 4 reports estimates for the likelihood of regaining coverage one month after separation, and the duration of uninsurance in months, respectively.[19] Column (3) reports the results from estimating a Cox Proportional Hazard model. Finally, columns (4)–(6) report results estimating the likelihood of moving to an employer plan, a Medicaid policy, or a marketplace plan, respectively, for those who find a new policy within 24 months after separation.

One month after separation, we estimate that unmarried women are 9.62 percentage points more likely to regain coverage, married male filers are 17.4 percentage points more likely to regain coverage, and married women are 23.56 percentage points more likely to regain coverage (measured as the combination of the coefficient on Female, Married, and Female x Married); differences by age and pre-separation earnings, on the other hand, are minimal. When accounting for the full 24-month post-separation observation window (col 3), we find that single women are 27.2% more likely to find coverage, and married men are 41.8% more likely to find coverage. Single women and married men also experience a shorter duration of uninsurance (1.876 months, and 2.638 months, respectively), and married women experience a 3.162 month shorter duration of uninsurance (measured as the combination of the coefficient on Female, Married, and Female x Married). Again, age and pre-separation wages do not meaningfully impact the duration of uninsurance.

[19] Appendix Table A5 reports results for the full population of policyholders who separate from an employer plan, regardless of whether they become unemployed.

While women and married policyholders generally exit uninsurance more quickly than men or unmarried policyholders, as seen in Table 2, these groups tend to take up different sources of coverage. This same pattern is seen in Table 4 holding all else equal. Married men are 15.2 percentage points more likely to purchase an employer plan (col 4), but only 1.18 and 0.732 percentage points (col 5, 6) more likely to move to a Medicaid or marketplace plan, respectively. Married women are 25.3 percentage points (col 4) more likely to be covered by an employer plan, compared with a 2.12 percentage point (col 5) reduction in the likelihood of moving to a Medicaid policy and a 1.48 percentage point increase (col 6) in the likelihood of moving to a marketplace plan. By contrast, single women are 12 percentage points more likely to move to a Medicaid policy (col 5) and 1.03 percentage points more likely to purchase a marketplace plan (col 6), but 0.656 percentage points *less likely* (col 4) to purchase an employer plan

4.4. Dynamics by Medicaid Expansion Status

The results in Table 4 also underscore that living in a Medicaid expansion state is strongly correlated with likelihood of finding new coverage, the duration of uninsurance, and the likelihood of moving to a Medicaid policy. In particular, individuals living in a Medicaid expansion state prior to their policy separation were 16.0 percentage points more likely to find coverage within a month (col 1) — and, more generally, were 36.6% more likely to find coverage when taking into account the full 24-month post-separation period (col 3). Moreover, those living in expansion states are 13.2 percentage points more likely to use Medicaid as a first source of new coverage (col 5). Taken together, these differences translate to shorter spells of uninsurance in expansion states lasting 50%, or 3 months, less than in non-expansion states (col 2).

In addition to these regressions, panels (b) and (c) of Figure 1 illustrate monthly sources of coverage following a loss for policyholders who had lived in an expansion state or a non-expansion state, respectively, at the time of their policy separation. Comparing across the two panels reveals several dramatic differences in post-separation coverage dynamics. First, the share of those in non-expansion states who transition to uninsurance one month after separation from an employer plan is 38% larger than in expansion states.[20] This gap grows to 50% when measured 24 months after separation;[21] in other words, individuals in expansion states are more likely to immediately find coverage and, conditional on transitioning to uninsurance, regain coverage more quickly. Second, employer plan coverage rates are similar across the two groups throughout the post-separation period,[22] suggesting that the labor market in these two groups of states may be similar enough to mitigate differential selection across labor markets. Finally, and most relevant, Medicaid plays an out-sized role as a source of coverage in expansion states.[23] In particular, Medicaid is *three times as likely* to be the source of coverage in every month throughout the 24-month post-separation observation period.

Finally, Table 5 reports the six-month transition rates across coverage types for Medicaid expansion and non-expansions states, respectively.[24] The persistence of uninsurance is 15% higher in non-expansion states compared to expansion states (51% vs. 44%, or a 7 percentage point difference). This difference is almost entirely driven by a 7 percentage point increase in the likelihood of transitioning from uninsurance to Medicaid in expansion states (10% vs.

[20] 42% of policyholders who separate are uncovered in expansion states, compared with 58% in non-expansion states.

[21] 18% of policyholders who separate are uncovered 24 months later in expansion states compared with 27% in non-expansion states.

[22] 36% are covered by an employer plan one month after separation in expansion states compared with 30% in non-expansion states. 24 months later, 61% are covered by an employer plan in expansion states compared with 59% in non-expansion states.

[23] A similar result is found for the full sample when looking at differences by Medicaid expansion and non-expansion states. These results are presented in Appendix Figure A1.

[24] The one-month transitions are given in Table A1 and reveal similar dynamics across expansion and non-expansion states.

3%). In addition, Medicaid coverage appears to be more stable in expansion states, with 61% of those with Medicaid coverage in month *m* having Medicaid coverage 6 months later, compared with just 55% in non-expansion states.

This section provides several related and reinforcing descriptive statistics that speak to the likely determinants of the duration of uninsurance spells following a separation from an employer policy. This analysis makes a compelling case for a causal effect of Medicaid expansion on post-separation coverage dynamics among individuals who are covered by an employer plan — a population that is, ex-ante, unlikely to be eligible for Medicaid. In the next section, we exploit a state-level experimental setting created by the Virginia expansion of Medicaid to study the causal effect the Medicaid expansion.

5. The Effects of Expanding Medicaid on Coverage Dynamics

Recall from Section 2 that the Medicaid Expansion is an optional provision of the ACA — that is, states can choose to expand Medicaid coverage to adults earning up to 138% FPL and receive federal funding to pay for 90–100% of the associated costs. When the ACA became effective in 2014, 25 states and the District of Columbia had chosen to expand Medicaid at the same time. Since then, 15 additional states have adopted the Medicaid expansion; these expansions offer a unique experimental setting to estimate the *causal* effect of increasing Medicaid access, holding fixed the other provisions of the ACA.

In this section, we focus specifically on the Virginia Medicaid expansion, which was implemented in January, 2019. Prior to the expansion, Medicaid access in Virginia had been severely limited: childless adults were ineligible, and eligibility among parents was restricted to those earning less than 38% of the Federal Poverty Level (FPL) equivalent to $7,896 for a parent in a family of three. Each month from 2017 to 2019, approximately 1,000,000 individuals were enrolled in Medicaid and CHIP.[25] Following Affordable Care Act (ACA) guidelines, Medicaid eligibility in Virginia was expanded to households earning below 138% of the FPL (e.g., $29,435 for a family of three). In the first month of the expansion, Medicaid enrollment surged by 17%, with steady growth thereafter, ultimately doubling to 2,000,000 enrollees by 2023.

Our data is well suited to leverage the expansion in Virginia to study the effect of expanding access to Medicaid for several reasons: first, the timing of this expansion allows us to observe plan separations prior to the policy change, providing the opportunity to identify a non-treated group; second, the timing of this expansion allows for post-separation observations that are uncontaminated by pandemic-era policies, such as expanded access to Medicaid and unemployment insurance; third, Virginia is a large and diverse state with its population that is spread across urban, suburban, and rural communities and comprises diverse political ideology and educational attainment.[26] Appendix Table A2 additionally provides selected demographic characteristics from the American Community Survey comparing Virginia's population to the overall U.S. in 2018. This table shows that Virginia has a similar age and sex composition to the U.S., though a higher median income, a higher share with a bachelors degree or above, and a lower share of uninsured.

[25]Data Source: Monthly State Medicaid Enrollment, Centers for Medicaid and Medicare Services.

[26]As a reminder, our analysis begins in 2016; we consider a natural bound of our analysis 2019, before the onset of the pandemic. Between 2015 and 2019, four states adopted Medicaid expansion: Montana (January 2016), Louisiana (July 2016), Virginia (January 2019), and Maine (January 2019). We exclude Montana because its January 2016 effective date does not allow us to observe plan separations prior to the Medicaid expansion. We exclude Maine because its Medicaid expansion was retroactive to mid-2018, making the pre and post period difficult to disentangle. We exclude Louisiana due to data issues—the 1095 data from Louisiana appear to be incomplete based on internal validation exercises. For these reasons, we focus solely on Virginia for this analysis.

For our empirical analysis, we will use a difference-in-difference estimation strategy that compares post-separation coverage dynamics for (1) individuals who separated from their employer health plan in Virginia before and after the Medicaid expansion to (2) individuals who separated from their employer health plan during these same month-years in all other non-expansion states.[27] This identification strategy will control for any differences between Virginia and other non-expansion states to the extent that these differences — including labor market differences, health differences, and other elements of the social safety net — are time-invariant during this short time period.[28]

5.1. Analysis Sample

Our analysis sample is defined by the months surrounding the January 1st, 2019 expansion. We focus on four cohorts: those who separated from their employer plan either in 2018 (one year before the Medicaid expansion) or in 2019 (the year of the Medicaid expansion) and who lived in either a non-expansion state or in Virginia at the time of their separation. To draw our data, we follow the procedures outlined in Section 3, with a few modifications to account for both the pandemic and the experimental setting. First, we focus on separations that occur between January and June of either 2018 (pre-treatment) or 2019 (post-treatment). Accordingly, we define pre-treatment cohorts to be policyholders who separate from an employer

[27] Our control group includes the following states: Alabama, Florida, Georgia, Kansas, Missouri, Mississippi, Nebraska, Oklahoma, South Carolina, South Dakota, Tennessee, Texas, Wisconsin, and Wyoming. Although Idaho and Utah did not expand Medicaid until Jan 1, 2020, each of these states was partially treated in 2019, so we exclude these states from our control group. Specifically, Idaho began expanded enrollment in November, 2019, and Utah implemented a bridge program to Medicaid expansion beginning in April, 2019. Finally, we exclude North Carolina because North Carolina implemented a major change to its Medicaid program, switching from a fee-for-service to a managed care network, where private companies, known as managed care organizations (MCOs), would be responsible for coordinating and delivering Medicaid services to beneficiaries. We note that our results are not sensitive to including Idaho, Maine, North Carolina, and Utah.

[28] To the best of our knowledge, there were no concurrent changes in access to other social safety net programs (unemployment insurance benefits, SNAP, WIC, TANF, etc) in Virginia that occurred at the same time as Medicaid expansion.

plan separations between January and June of 2018. We then limit our post-separation observations to the last six months *in the same year as separation*. This truncation ensures that the our outcomes for the pre-treatment group are not affected by the treatment itself (the January 2019 Virginia Medicaid Expansion). Likewise, our outcomes for the post-treatment group will avoid any confounding effects of the pandemic and associated policy.[29] Moreover, by comparing pre- and post-treatment separations that occur in the same month, one year apart, we hope to hold fixed, at a minimum, the influence of monthly employment dynamics.

Second, we take a slightly more restrictive definition of individuals who drop coverage due to unemployment. In particular, we limit our analysis to those who received UI *in the same year as plan separation*. We do this because pandemic-era policies dramatically expanded access to the unemployment insurance system beginning in March, 2020. This is a departure from our previous analysis sample, which included individuals who received UI in both the year of plan separation and the following year.

Table 6 reports summary statistics for this analysis sample. Columns (1) and (2) describe pre-treatment policyholders, or those who separated from their plan between January and June of 2018 and were living in Virginia or a non-expansion state, respectively. Similarly, columns (3) and (4) describes post-treatment policyholders, or those who separated from their plan between January and June of 2019 and were living in Virginia or a non-expansion state, respectively. Panel A summarizes demographic characteristics in the year before the plan separation. Panels B and C summarize outcome variables and preview our results. In general, policyholders living in Virginia are more likely to be female, are less likely to be joint-filers, are slightly

[29] Uninsurance was especially risky during this time period: the onset of the pandemic imposed a broad increase in ex-ante expected health care expenditures for the U.S. population due to the perceived risk of infection and hospitalization simultaneously with increased macroeconomic turmoil and reduced financial ability to weather large medical expenditures. In light of this, the federal government undertook an historic expansion of the social safety net through increased emergency funds, emergency Medicaid waivers that allowed states to streamline enrollment processes, expand coverage options, increase access to telehealth services, and relax typical eligibility verification to permit continuous enrollment in Medicaid. In addition, the Federal government expanded access to Unemployment Insurance to the self-employed and broadly expanded benefits.

older, and earn higher wages. Importantly in the difference-in-differences context, differences in these characteristics appear to be roughly stable in 2018 and 2019.

As a preview of the results, when comparing mean changes between Virginia and other non-expansion states in the pre- and post-treatment cohort (columns 1–4), the gap between Virginia policyholders and those living in non-expansion states widens for our three primary outcomes (Panel B). Policyholders who separate from an employer plan in Virginia after the Medicaid expansion are more likely to be covered one month after separation and are less likely to remain uncovered through the end of 2019. In addition, they experience a shorter spell of uninsurance when compared to these differences for the pre-treatment cohort. Finally, differences in the first source of coverage following a plan separation, particularly for Medicaid, widen substantially (Panel C) after the expansion: Virginians are 12 percentage points more likely to move to Medicaid compared to 3 percentage points less likely prior to the expansion.

5.2. Empirical Strategy: Difference-in-Differences

We estimate the following difference-in-differences equation:

$$Y_i = \beta_0 + \beta_1\, Treat_i + \beta_2\, Post_i + \beta_3\, Treat_i \times Post_i + \Gamma_i + \varepsilon_{it}, \tag{1}$$

where $Treat_i$ is a dummy variable identifying separated policyholders living in Virginia, and $Post_i$ is a dummy variable identifying policyholders who separate from their plan in 2019, after Virginia's expansion of Medicaid. The vector Γ_i includes a series of fixed effects identifying gender, joint-filing, and month of separation, as well as continuous controls for age and wages earned in the year before separation. Y_i represents our outcomes of interest: an indicator that coverage is regained in the month after separation, the total number of observed months of uninsurance through December, or an indicator for the first source of coverage (Medicaid, employer plan, or Exchange). We study the first source of coverage to confirm the mechanism of the Medicaid expansion: if reduced duration of uninsurance is caused by the Medicaid expansion, we would expect to see an increase in the likelihood that Medicaid serves as the first source of coverage. The coefficient, $\hat{\beta}_3$, captures the effect of expanded Medicaid access in Virginia on outcome Y_i relative to a control group of separated policyholders living non-expansion states.

To study the effect of Medicaid access on the instantaneous likelihood of finding new coverage, we must account for the truncation of the outcome variable. To do so, we incorporate a Cox Proportional Hazard Model into a difference-in-differences[30] environment as follows:

$$h(m|x_j) = h_0(m)\exp(\beta_1\, Treat_i + \beta_2\, Post_i + \beta_3\, Treat_i \times Post_i + \Gamma_i + \varepsilon_{it}). \tag{2}$$

In this model, $\hat{\beta}_3$ captures any change in the hazard ratio that occurs after the Medicaid expansion. As before, the vector Γ_i controls for gender, joint-filing, month of separation, age, and wages from the year before separation.

5.3. Results

We report the estimates of the effect of the Virginia Medicaid expansion in Table 7. Columns (1) – (3) provide estimates of the effect of the Medicaid expansion on the likelihood of finding insurance one month after separation, the duration of uninsurance, and the hazard ratio. Columns (4) – (6) report the likelihood that the first source for those who find coverage by December is either a new employer plan, Medicaid, or an Exchange policy, respectively. In

[30]Similar methodology is used in Mastrobani and Pinotti (2015); this analysis relies on a Cox difference-in differences model to estimate the impact of legal status on recidivism.

addition, we report the pre-treatment mean for each outcome, or the unconditional mean for policyholders who separate from a plan in Virginia in 2018. In all cases, robust standard errors are reported in parentheses.

5.3.1. Baseline Results

We estimate that the Virginia Medicaid expansion increased the likelihood of finding coverage one month after separation by 16% (6.60 percentage points/40.8%, panel A col 1). In addition, the Medicaid expansion reduced the duration of uninsurance by 12% (-0.487/4.093 months, panel A col 2). Finally, the instantaneous likelihood of regaining coverage increased by 16.6% (panel A col 3).

To provide additional support for identification, we split our results based on wage earnings in the year prior to separation; lower income households are more likely to be treated than higher income households. Panels B and C of Table 7 reports estimates for individuals who were in the bottom (top) quartile of wage distribution prior to plan separation. We find that low wage workers are more than three times more likely to find coverage one month later (11.6 percentage points compared with 3.76 percentage points, panels B and C col 1). More generally, low wage workers are 33.3% more likely to find coverage (panel B col 3); by comparison, we do not find a statically significant effect on the instantaneous likelihood of finding coverage for high wage workers.

We also engage in a falsification exercise where we estimate the difference-in-difference model as though each of the control states were treated, using all other states in our analysis as the control group. These results are shown in Figure 2. For reference, we also show baseline estimates of the effect of the Virginia Medicaid expansion in this figure.

In panel (a) we see that in all other control states except for Missouri, the change in the likelihood of finding coverage one month after plan separation is small and statistically insignificant. While Missouri's estimate is statistically significant, it is wrong-signed relative to the effect of Medicaid (a small reduction in the likelihood of finding coverage). Further investigation reveals that the topic of Medicaid expansion became a focus of significant political and public interest in 2019, culminating in a 2020 ballot initiative to force the legislature to expand Medicaid that was approved. With this in mind, it is plausible that the attention and scrutiny of the state's Medicaid program that was brewing in 2019 affected Medicaid enrollment.

In panels (b) and (c) we see that Virginia was the only state that experienced a decrease in the duration of uninsurance paired with an increase in the likelihood of finding coverage. Most of our control states saw no meaningful or statistically significant change in either outcome. At the same time, a few states stand out with opposite-signed estimates. Two estimates from small states (South Dakota and Wyoming) are only marginally statistically significant; to the best of our knowledge, nothing about either state's Medicaid program or the political and public discourse surrounding Medicaid expansion changed. Tennessee, however, proposed a major modification to their Medicaid program in 2019 by applying for a waiver from the ACA requirements to implement a modified version of Medicaid expansion through a program called "TennCare III." Although this waiver was not ultimately approved until April, 2022, the effort to launch TennCareIII suggests that the state's Medicaid program was a focal point, and this may have affected Medicaid enrollment.

Next, we study the effect of the Medicaid expansion on the first source of coverage among those individuals who regain coverage by the end of the year. If the Medicaid expansion causes changes in coverage dynamics, then we would expect that shortened uninsurance spells and increased coverage should coincide with an increase in Medicaid as a first source of .

Consistent with this mechanism, we find a 15.1 percentage point increase in the likelihood that Medicaid serves as the first source of coverage — a 154% increase relative to the pre-treatment mean (panel A col 5). Low wage workers are nearly four times more likely to move to a Medicaid plan than high wage workers (26 percentage points compared with 6.84 percentage points, panels B and C col 5). At the same time, we estimate a 10.1 and a 3.56 percentage point reduction in the likelihood that the first source of coverage is an employer plan or an exchange policy, respectively. While these estimates suggest that individuals shift from other coverage sources to Medicaid as a first source of coverage, they do not rule out that some Medicaid recipients also came from a counterfactual state of uninsurance.

5.3.2. Heterogeneity by Gender

Generally, women are more likely to have insurance relative to men, both before and after the enactment of the ACA. This is likely driven, in part, by the fact that women are more likely than men to qualify for Medicaid because they tend to have lower incomes and are more likely to belong to one of Medicaid's adult eligibility categories (pregnant, parent of a dependent child, senior, or person with a disability). In addition, women may have a stronger preference for health insurance coverage compared to men (Gomez et al., 2022). Consistent with this, the descriptive evidence in Table 4 shows that women are more likely to regain coverage and experience a shorter duration of uninsurance than men. In Figure 3, we explore how access to Medicaid affects these dynamics differently by gender. Point estimates are reported in Appendix Table A7.

Overall, we find that the effect of the Medicaid expansion was stronger for women than for men. The likelihood of regaining coverage one month after separation increased by 8.55 percentage points (21%) for women, compared to 4.78 percentage points (12%) for men. Women's duration of uninsurance decreased by 0.623 months (16%) compared with 0.357

months (9%) for men. Finally, using a hazard model we estimate that women are 21.1% more likely to regain coverage, compared with a 12.7% increase for men.

Turning to results on the first source of coverage, both women and men experience a large increase in the likelihood that Medicaid serves as a first-source of coverage (17.2 percentage points for women and 13.1 percentage points for men). However, the effect relative to the pre-treatment mean is larger for men (300%) than for women (110%). This larger relative likelihood for men is entirely consistent with the mechanism of the Medicaid expansion: men were less likely to qualify for Medicaid prior to ACA expansion.

Both female and male policyholders exhibit some shifting from employer plans and exchange policies to Medicaid. They are similarly likely to shift away from employer plans (10.8 percentage point decrease for women and 9.32 percentage point decrease for men), but women are more likely to shift away from an exchange plan (5.29 percentage points, or a 54% reduction) than men (1.90 percentage points or a 29% reduction).

5.3.3. Heterogeneity by Marital Status

Next, we examine heterogeneity among individuals based on marital status. Marriage offers an additional channel for health insurance access through dependent coverage on a spouse's plan. As a reminder, our analysis focuses on *policyholders* who separate from their own employer plan. However, if a spouse loses their health insurance, this is considered to be a "qualifying life event," enabling a special enrollment period where the uninsured spouse may be added as a dependent on their partner's plan. This exception applies regardless of whether the spouse's separation from their own employer plan was voluntary or involuntary. Given that marriage provides an additional avenue for gaining coverage, we would expect that those who are married would be less impacted by the Medicaid expansion. With this in mind, Figure 3 reports estimates for unmarried and married individuals.

Policyholders who were unmarried are 25% (8.36 percentage points/33.4%) more likely to gain coverage in the month after separation due to the Medicaid expansion, while married policyholders were 6% more likely (3.17 percentage points/54.5%). Uninsurance spells declined for both groups, but declined more so for unmarried policyholders (14% vs. 5%). More generally, the instantaneous likelihood of regaining coverage is nearly five times larger for unmarried and married policyholders (24.5% and 5.1%, respectively); notably, the estimate for married policyholders is not statistically significant at any conventional levels.

Both married and unmarried policyholders exhibit some shifting from employer plans and exchange policies to Medicaid. The shifts are more pronounced for unmarried policyholders, who experience declines in employer coverage of 12.6 percentage points compared to the 4.93 percentage points experienced by married policyholders. Similarly, exchange policies as a first source of coverage declined by 4.47 percentage points and 2.00 percentage points respectively for unmarried and married individuals. In general, there appears to be less crowd-out of Medicaid as a first source of coverage for those who are married compared to those who are unmarried.

5.4. Discussion

Most Americans rely on employer health plans for coverage, making them vulnerable to spells of uninsurance when faced with job loss. Our analysis indicates that the ACA's Medicaid expansion provides a path to coverage stability by increasing the likelihood of finding coverage and reducing the duration of uninsurance for for individuals experiencing job loss and subsequent loss of health insurance. Additionally, our findings demonstrate that the ACA broadened the scope of the Medicaid safety net beyond the chronically low-income to include those who previously held stable, high-quality jobs.

Spells of uninsurance increase the cost of health care and, therefore, reduce health care consumption. Gai and Jones (2020) show that insurance instability is associated with forgone preventative and other medical services. Roberts and Pollack (2016) find that churn is associated with sub-optimal timing of care as evidenced through increased emergency room visits and inpatient hospital stays. Insurance instability is particularly harmful for those with chronic conditions who require access to ongoing care. Kressin et al. (2020) find that insurance instability was associated with worse health outcomes, particularly for Black and Hispanic patients, as measured by higher rates of uncontrolled blood pressure.

Spells of uninsurance are also associated threats to financial risk. In particular, costly treatment may be unavoidable for those facing acute health shocks, thereby increasing the risk of financial setbacks during periods of uninsurance. Medicaid is shown to mitigate this risk. For example, Gross and Notowidigdo (2011) finds that pre-ACA Medicaid expansions reduced the likelihood of personal bankruptcy. More broadly, Hu et al. (2018); Kuroki (2021); Miller et al. (2021) find evidence that the ACA's Medicaid expansion was associated with improved financial well-being.

Though beyond the scope of this paper, our work compliments and expands literature on the negative effects of coverage stability by taking a step back and studying how Medicaid can affect coverage stability, itself. In addition, our work has implications for proposals that incorporate work requirements into Medicaid eligibility rules. Recently, a number of states have debated modifications to their state Medicaid programs that would condition coverage on work and reporting requirements, similar to public assistance programs like Temporary Assistance for Needy Families (TANF). Sommers et al. (2020) finds that Medicaid work requirements in Arkansas in 2018 led to a decrease in Medicaid or Marketplace coverage of 13.2 percentage points and an increase in the uninsurance rate of 7.1 percentage points among Arkansans aged 30-49. These results suggest that the imposition of work requirements associated with Medicaid, especially in Medicaid expansion states, is likely to decrease the effectiveness of the ACA's Medicaid expansion in reducing coverage instability.

6. Conclusion

This study leverages the new tax reporting requirements mandated by the ACA, offering a more comprehensive understanding of coverage dynamics pertinent to the majority of the population. We present novel descriptive statistics to characterize coverage dynamics for those who separate their employer health coverage. Among policyholders who experience both separation from an employer plan and unemployment, we provide insights into the average duration of uninsurance, the sources of coverage during this period, and detailed transitions across various types of coverage.

Previous research examining coverage instability and the mitigating effects of Medicaid primarily focuses on the chronically low-income demographic. In contrast, we demonstrate that the 60% of individuals in the U.S. covered by employer plans are also vulnerable to coverage instability, due in part to labor market dynamics.

We offer the first causal estimates of the ACA's Medicaid expansion's impact on high-frequency coverage dynamics by leveraging a compelling source of variation — the 2019 Virginia Medicaid expansion. Our findings reveal that the state-level Medicaid expansion in Virginia increased the likelihood of regaining coverage by 17% for Virginians who lost their employer plan. We also show that these effects were the strongest for female and unmarried policyholders.

Moreover, the expansion induced a shift in the composition of the initial source of new coverage, with Medicaid experiencing a 15.1 percentage point increase and private coverage witnessing a 13.7 percentage point decrease. This pattern suggests some crowd-out effect of private insurance by Medicaid as a first source of coverage, which would lower the return on investment of the Medicaid expansion in terms of general social welfare. At the same time, expanded access to Medicaid also likely allows for better matching in the labor search process. For instance, individuals may now have the flexibility to choose more suitable employment opportunities rather than settling for less ideal positions solely to secure health insurance coverage. Although beyond the scope of this paper, future research may investigate whether the expanded access to Medicaid leads to higher job match quality and higher earnings.

References

Agarwal, Sumit D, and Benjamin D Sommers. 2020. "Insurance coverage after job loss—the importance of the ACA during the Covid-associated recession." *New England Journal of Medicine*, 383(17): 1603–1606.

Bertakis, KD, R Azari, LJ Helms, EJ Callahan, and JA Robbins. 2000. "Gender differences in the utilization of health care services." *Journal Family Practice*, 49(2): 147–152.

Corallo, Bradley, Rachel Garfield, Jennifer Tolbert, and Robin Rudowitz. 2021. "Medicaid Enrollment Churn and Implications for Continuous Coverage Policies." Kaiser Family Foundation.

Cutler, David M., and Alexander M. Gelber. 2009. "Changes in the Incidence and Duration of Periods without Insurance." *New England Journal of Medicine*, 360(17): 1740–1748.

East, Chloe, and David Simon. 2022. "How Well Insured are Job Losers? Efficacy of the Public Safety Net."

Einav, Liran, and Amy Finkelstein. 2023. "The Risk of Losing Health Insurance in teh United States is Large, and Remained So After the Affordable Care Act." *Proceedings of the National Academy of Sciences*, 120(18): 1–7.

Fairlie, Robert W, and Rebecca A London. 2008. "The Dynamics of Health Insurance Coverage: Identifying Trigger Events for Insurance Loss and Gain." *Health Services and Outcomes Research Methodology*, 8(3): 159–185.

Frean, Molly, Jonathan Gruber, and Benjamin D. Sommers. 2017. "Premium Subsidies, the Mandate, and Medicaid Expansion: Coverage Effects of the Affordable Care Act." *Journal of Health Economics*, 53: 72–86.

Gai, Yunwei, and Kent Jones. 2020. "Insurance Patterns and Instability from 2006 to 2016." *BMC Health Services Research*, 20(1): 1–12.

Gomez, Ivette, Usha Ranji, Alina Salganicoff, and Brittni Frederiksen. 2022. "Medicaid Coverage for Women." Kaiser Family Foundation.

Graves, John A, and Sayeh S Nikpay. 2017. "The Changing Dynamics Of US Health Insurance And Implications For The Future Of The Affordable Care Act." *Health Affairs*, 36(2): 297–305.

Gross, Tal, and Matthew J. Notowidigdo. 2011. "Health insurance and the consumer bankruptcy decision: Evidence from expansions of Medicaid." *Journal of Public Economics*, 95(7-8): 767–778.

Guevara, James P, Jeanhee Moon, Elizabeth M Hines, Ettya Fremont, Angie Wong, Christopher B Forrest, Jeffrey H Silber, and Susmita Pati. 2014. "Continuity of Public Insurance Coverage: A Systematic Review of the Literature." *Medical Care Research and Review*, 71(2): 115–137.

Ham, John C, Xianghong Li, and Lara Shore-Sheppard. 2009. "Seam bias, multiple-state, multiple-spell duration models and the employment dynamics of disadvantaged women." National Bureau of Economic Research.

Hu, Luojia, Robert Kaestner, Bhashkar Mazumder, Sarah Miller, and Ashley Wong. 2018. "The Effect of the Affordable Care Act Medicaid Expansions on Financial Wellbeing." *Journal of Public Economics*, 163: 99–112.

Kressin, Nancy R., Norma Terrin, Amresh D. Hanchate, Lori Lyn Price, Alejandro Moreno-Koehler, Amy Leclair, Jillian Suzukida, Sucharita Kher, and Karen M. Freund. 2020. "Is Insurance Instability Associated with Hypertension Outcomes and Does This Vary by Race/Ethnicity?" *BMC Health Services Research*, 20(1): 1–8.

Kuroki, Masanori. 2021. "The effect of health insurance coverage on personal bankruptcy: evidence from the Medicaid expansion." *Review of Economics of the Household*, 19(2): 429–451.

Lurie, Ithai Z, and James Pearce. 2021. "Health Insurance Coverage in Tax and Survey Data." *American Journal of Health Economics*, 7(2): 164–184.

Mastrobani, Giovanni, and Paolo Pinotti. 2015. "Legal Status and the Criminal Activity of Immigrants." *American Economic Journal: Applied Economics*, 7(2): 175–206.

Miller, Sarah, Luojia Hu, Robert Kaestner, Bhashkar Mazumder, and Ashley Wong. 2021. "The ACA Medicaid Expansion in Michigan and Financial Health." *Journal of Policy Analysis and Management*, 40(2): 348–375.

Proctor, Bernadette D, Jessica L Semega, and Melissa A Kollar. 2016. "Income and Poverty in the United States: 2015 Current Population Reports." US Census Bureau.

Roberts, Eric T., and Craig Evan Pollack. 2016. "Does Churning in Medicaid Affect Health Care Use?" *Medical Care*, 54(5): 483.

Schaller, Jessamyn, and Ann Huff Stevens. 2015. "Short-run effects of job loss on health conditions, health insurance, and health care utilization." *Journal of Health Economics*, 43: 190–203.

Schaller, Jessamyn, and Mariana Zerpa. 2019. "Short-run effects of parental job loss on child health." *American Journal of Health Economics*, 5(1): 8–41.

Sommers, Benjamin D, Lucy Chen, Robert J Blendon, E John Orav, and Arnold M Epstein. 2020. "Medicaid Work Requirements In Arkansas: Two-Year Impacts On Coverage, Employment, And Affordability Of Care: Study examines the impact of the Arkansas Medicaid work requirement before and after a federal judge put the policy on hold." *Health Affairs*, 39(9): 1522–1530.

Swartz, Katherine, and Timothy D McBride. 1990. "Spells Without Health Insurance: Distributions of Durations and Their Link to Point-in-Time Estimates of the Uninsured." *Inquiry*, 27(3): 281–288.

Swartz, Katherine, John Marcotte, and Timothy D McBride. 1993. "Personal Characteristics and Spells Without Health Insurance." *Inquiry*, 30(1): 64–76.

Vistnes, Jessica P, and Joel W Cohen. 2018. "Duration Of Uninsured Spells For Nonelderly Adults Declined After 2014." *Health Affairs*, 37(6): 951–955.

Table 1
Summary Statistics

	Population	Unemployed		
		All States	Expansion	Non-Expansion
	(1)	(2)	(3)	(4)
Female	0.4576	0.418	0.418	0.418
Married	0.425	0.378	0.378	0.379
Age				
18–25	0.0904	0.0628	0.0643	0.0599
26–44	0.549	0.518	0.525	0.505
46–62	0.360	0.419	0.411	0.435
Wages	62,093	55,673	56,784	53,453
Unemployed	0.153	.	.	.
Employed in 2017	0.912	0.888	0.887	0.889
2015 MAGI				
$<$ 100% FPL	0.0488	0.0491	0.0479	0.0515
100–138%	0.0509	0.0583	0.0537	0.0673
138–400%	0.469	0.507	0.497	0.527
$>$ 400% FPL	0.431	0.386	0.401	0.354
Covered at t=1	0.659	0.523	0.576	0.416
Months of Uninsurance	3.691	4.887	4.085	6.492
Always Uninsured	0.051	0.063	0.043	0.102
N	10,643,393	1,633,155	1,088,857	544,298

Notes: This table summarizes the population of interest, policyholders aged 18–62 who separate from an employer plan in 2016 (col 1), in addition to our analysis subsample of those who also receive unemployment benefits (cols 2–4). Modified Adjusted Gross Income relative to the Federal Poverty Limit (FPL) is based on 2015 income and household size. Marital status is determined by whether the household filed jointly in 2015. The share of unemployed individuals in columns 2–4 are identified as those who receiving income reported on Form 1099-G in either 2016 or 2017. Coverage and duration information is based on monthly coverage reported on Forms 1095-A, -B, and -C and is observed in our data for 24 months after initial policy separation.

Table 2
Summary Statistics Based on First Month of Coverage:
Policyholders Who Separate From an Employer Plan in 2016 and Become Unemployed

	First Month of Coverage					
	1	3	6	12	24	Never Covered
Female	0.447	0.417	0.393	0.360	0.325	0.313
Joint-Filer	0.442	0.355	0.300	0.269	0.227	0.235
Age						
18–25	0.0761	0.0457	0.0499	0.0516	0.0639	0.0539
25–44	0.507	0.534	0.531	0.545	0.538	0.512
45–62	0.417	0.420	0.419	0.403	0.398	0.434
2015 Wages	\$57,208	\$57,595	\$54,073	\$52,415	\$47,118	\$46,192
Employed in 2017	0.894	0.915	0.916	0.912	0.822	0.714
2015 MAGI						
$<$100% FPL	0.0655	0.0262	0.0303	0.0334	0.0408	0.0424
100–138% FPL	0.0655	0.0419	0.0469	0.0502	0.0623	0.0697
138–400% FPL	0.461	0.529	0.554	0.577	0.610	0.619
$>$ 400% FPL	0.408	0.403	0.368	0.339	0.287	0.268
First Source of Coverage						
Employer Plan	0.646	0.673	0.725	0.740	0.727	-
Marketplace Plan	0.0512	0.161	0.0846	0.0810	0.0825	-
Medicaid Plan	0.279	0.136	0.174	0.160	0.162	-
N	855,692	105,881	44,416	18,051	4,898	102,986

Notes: This table summarizes policyholders based on the first month that we observe new coverage after having separated from an employer plan in 2016. The sample is limited to those who receive unemployment income in 2016 or 2017. See also Table 1 notes.

Table 3
Summary Statistics Based on First Source of Coverage:
Policyholders Who Separate From an Employer Plan in 2016 and Become Unemployed

	First Source of Coverage		
	Employer	Medicaid	Marketplace
Female	0.398	0.526	0.484
Joint Filer	0.433	0.244	0.407
Aged 18–25	0.0680	0.0762	0.0192
Aged 26–44	0.512	0.581	0.409
Aged 45–62	0.420	0.343	0.571
2015 Wages	61,590	37,588	59,646
2015 MAGI			
$<$100% FPL	0.0269	0.136	0.0242
100–138% FPL	0.0385	0.132	0.0351
138–400% FPL	0.469	0.587	0.519
$>$ 400% FPL	0.465	0.145	0.422
Employed, 2017	0.938	0.837	0.804
Months of Uninsurance	3.820	2.927	3.895
N	1,020,457	351,491	124,668

Notes: This table summarizes policyholders who receive unemployment income in 2016 of 2017 based on their first source of coverage after a separation from an employer plan in 2016. See also Table 1 notes.

Table 4
Post-Separation Coverage Dynamics:
Policyholders Who Separate From an Employer Plan in 2016

	Coverage Dynamics			First Source of Coverage		
	Covered at t=1	Months of Uninsurance	Hazard Ratio	Employer Plan	Medicaid Plan	Marketplace Plan
	(1)	(2)	(3)	(4)	(5)	(6)
Female	0.0962***	-1.876***	1.272***	-0.00656***	0.120***	0.0103***
	(0.000978)	(0.0137)	(0.00265)	(0.000949)	(0.000818)	(0.000323)
Married	0.174***	-2.638***	1.418***	0.152***	0.0118***	0.00732***
	(0.00104)	(0.0145)	(0.00314)	(0.00122)	(0.000982)	(0.000353)
Female *x* Married	-0.0346***	1.352***	0.850***	0.108***	-0.153***	-0.00285***
	(0.00161)	(0.0200)	(0.00293)	(0.00168)	(0.00119)	(0.000604)
Age	-0.00185***	0.0238***	0.997***	-0.00155***	-0.00195***	0.000969***
	(0.0000359)	(0.000492)	(0.0000762)	(0.0000419)	(0.0000357)	(0.0000133)
Wages	0.0000807***	-0.00417***	1.000***	0.000704***	-0.000743***	-1.48e-08
	(0.00000657)	(0.000187)	(0.00000761)	(0.0000294)	(0.0000318)	(0.00000218)
Expansion State	0.160***	-2.392***	1.366***	0.0389***	0.132***	-0.00411***
	(0.000810)	(0.0117)	(0.00240)	(0.000811)	(0.000530)	(0.000296)
Unconditional Mean	0.524	4.877	–	0.360	0.157	0.029
Month Fixed Effects	✓	✓	✓	✓	✓	✓
N	1,633,155	1,633,155	1,633,155	1,530,371	1,530,371	1,530,371

Notes: This table reports results from estimating a OLS model (cols 1,2, 4, 5, 6) and a Cox Proportional Hazard model (col 3). Analysis is based on policyholders who separate from an employer plan in 2016 and become unemployed in 2016 or 2017. Post separation monthly coverage is observed for 24 months for all individuals in this analysis. All specifications include monthly fixed effects identifying the month of separation in 2016. See Section 3 for more details.

Table 5
Six-Month Coverage Transitions

	$m+6$			
m	Uncovered	Employer	Marketplace	Medicaid
Panel A. All Unemployed				
Uncovered	47%	43%	3%	7%
Employer	12%	82%	2%	4%
Marketplace	8%	41%	46%	5%
Medicaid	7%	31%	1%	61%
Panel B. Expansion				
Uncovered	44%	44%	3%	10%
Employer	11%	83%	2%	5%
Marketplace	7%	41%	45%	7%
Medicaid	7%	31%	1%	61%
Panel B. Non-Expansion				
Uncovered	51%	43%	3%	3%
Employer	16%	81%	2%	2%
Marketplace	10%	40%	48%	2%
Medicaid	12%	32%	1%	55%

Notes: This table describes transitions for policyholders who separate from an employer plan in 2016 and also become unemployed in 2016 or 2017. These statistics reflect the likelihood of transitioning across coverage sources from month m to month $m+6$ for the full sample (Panel A), and separately for those living in expansion states (Panel B) and non-expansion states (Panel C) prior to their initial dropped coverage.

Table 6
Summary Statistics: VA Medicaid Expansion

	Pre-Treatment Cohort 2018 Separations		Post-Treatment Cohort 2019 Separations	
	VA	Non-Exp	VA	Non-Exp
	(1)	(2)	(3)	(4)
Panel A. Prior Year Characteristics				
Female	0.476	0.472	0.486	0.461
Joint Filer	0.348	0.346	0.333	0.342
Age	43.59	43.27	43.6	43.45
Wages	59,207	54,059	59,748	56,735
N	12,710	170,902	11,025	161,984
Panel B. Post-Separation Coverage Outcomes				
Covered at t=1	0.407	0.353	0.467	0.348
Months of Uninsurance	4.097	4.476	3.634	4.499
Always Uninsured	0.276	0.324	0.219	0.330
N	12,710	170,902	11,025	161,984
Panel C. First Source of Coverage Following Plan Separation				
Employer Plan	0.794	0.773	0.700	0.782
Medicaid Plan	0.0935	0.123	0.238	0.119
Exchange Plan	0.0739	0.0685	0.0429	0.0664
N	10,424	131,010	9,704	122,004

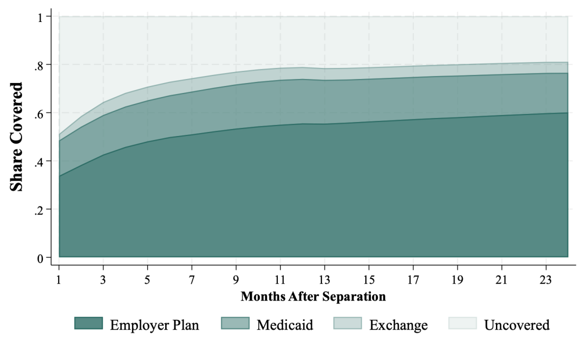
Notes: This table summarizes our primary analysis sample for our difference-in-differences analysis. Treated individuals are those who separate from an employer plan after Medicaid expansion in Virginia. Control individuals separate during these same months states that had not expanded Medicaid by the end of 2018. Post-separation coverage observed through the end of the same calendar year as separation. Panel A describes individual characteristics, measured in the year prior to separation. Panel B describes post-separation coverage outcomes, where the post-separation period is observed through the end of each cohort-specific calendar year. Panel C describes the first source of coverage for individuals who re-gain coverage within the calendar year. See also Table 1 for more details.

Table 7
Coverage Dynamics: Baseline Estimate of the Effect of ACA's Medicaid Expansion

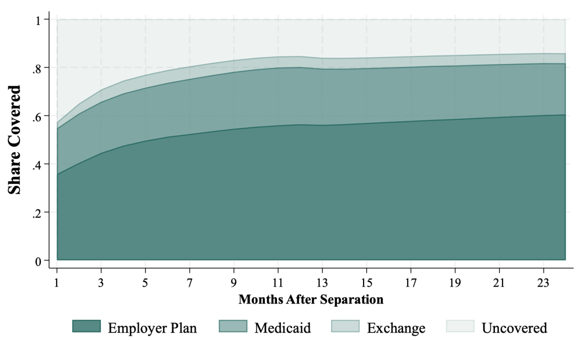
	Coverage Dynamics			First Source of Coverage		
	Covered at t=1	Months of Uninsurance	Hazard Ratio	Employer Plan	Medicaid Plan	Marketplace Plan
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Baseline Result						
2019 × VA	0.0660***	-0.487***	1.166***	-0.101***	0.151***	-0.0356***
	(0.00635)	(0.0440)	(0.0182)	(0.00661)	(0.00508)	(0.00403)
Pre-Treatment Mean	0.408	4.093	–	0.777	0.0980	0.0814
N	356,621	356,621	356,621	237,114	237,114	237,114
Panel B. Low Wages						
2019 × VA	0.116***	-0.982***	1.333***	-0.178***	0.260***	-0.0586***
	(0.0135)	(0.0954)	(0.0441)	(0.0152)	(0.0133)	(0.00877)
Pre-Treatment Mean	0.405	4.356	–	0.663	0.200	0.0902
N	84,446	84,446	84,446	59,052	59,052	59,052
Panel C. High Wages						
2019 ×	0.0376**	-0.182*	1.050	-0.0309**	0.0684***	-0.0254***
	(0.0126)	(0.0819)	(0.0304)	(0.0106)	(0.00605)	(0.00720)
Pre-Treatment Mean	0.493	3.390	–	0.868	0.025	0.073
N	81,463	81,463	81,463	54,335	54,335	54,335

Notes: This table reports the effect of Medicaid Expansion on post-separation coverage dynamics using a difference-in-differences identification strategy. Panel A reports the baseline results, while Panels B and C report estimates respectively for the bottom 25th percentile and top 25th percentile of the wage distribution where wages are measured in the year before losing coverage. Pre-treatment mean measured in Virginia in 2018. See Section 5 for more details.

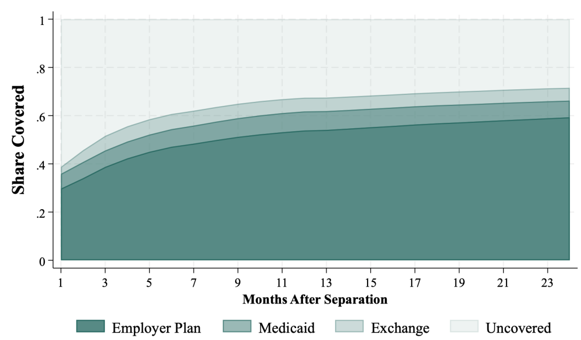
Figure 1. Post-Separation Monthly Coverage



(a) All



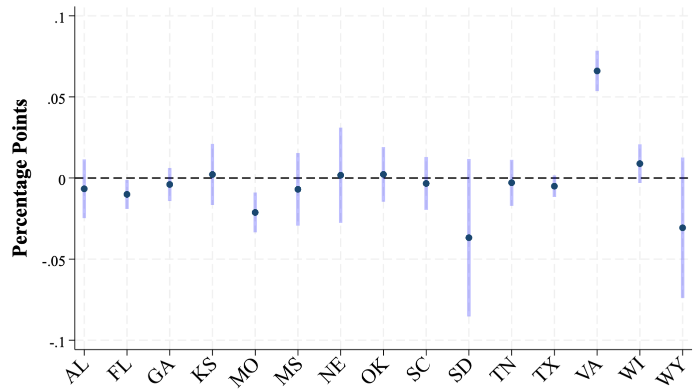
(b) Medicaid Expansion



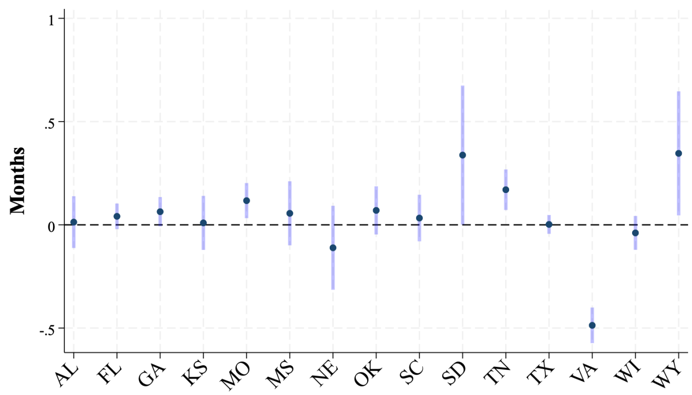
(c) Medicaid Non-Expansion

Notes These figures depict the breakdown of the source of health insurance coverage in each of the 24 months after a policyholder who becomes unemployed separates from an employer plan in 2016 based on whether the policyholder lived in an expansion state (panel a) or a non-expansion state (panel b).

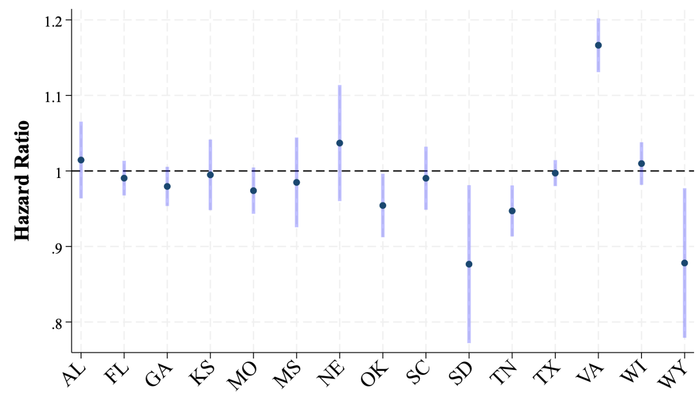
Figure 2. Falsification Tests



(a) Probability of Finding Coverage in $t = 1$



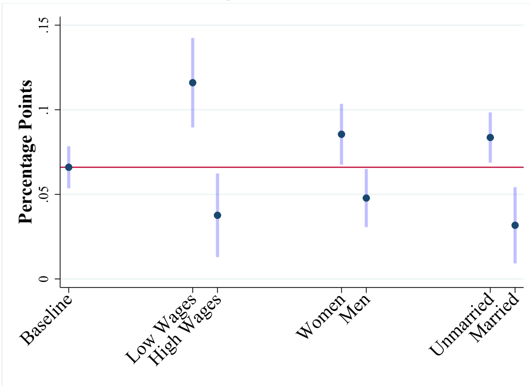
(b) Duration of Uninsurance



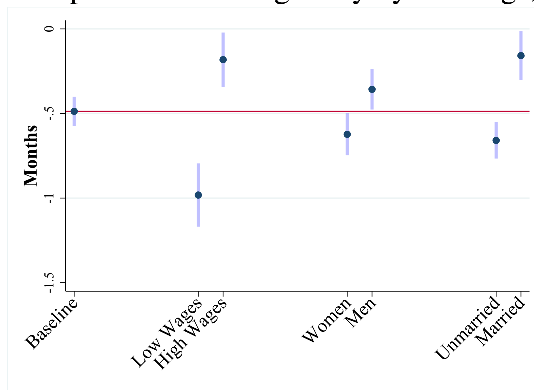
(c) Instantaneous Likelihood of Finding Coverage

Notes These figures depict estimates of a placebo difference-in-difference, where each control state is designated as a treated state. Each graph also depicts our baseline estimate of the effect of the Virginia Medicaid expansion. Panel (a) reports estimates and 95% confidence intervals for the outcome measuring the likelihood of finding coverage one month after plan separation, panel (b) reports estimates for the duration of uninsurance, measured in months, and panel (c) reports estimates of the hazard ratio, or the instantaneous likelihood of finding coverage after plan separation.

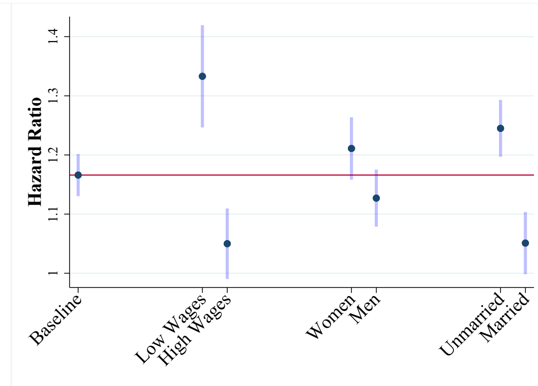
Figure 3. Effect of Medicaid Expansion: Heterogeneity by Earnings, Gender, and Marital Status



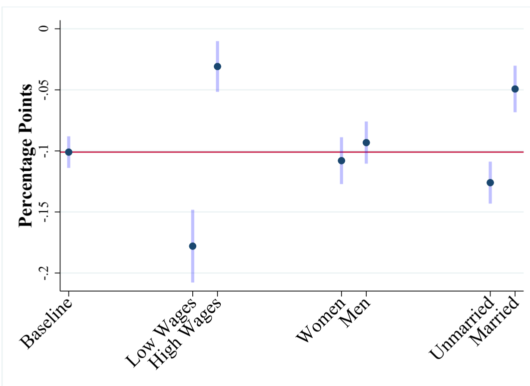
(a) Probability of Finding Coverage in $t = 1$



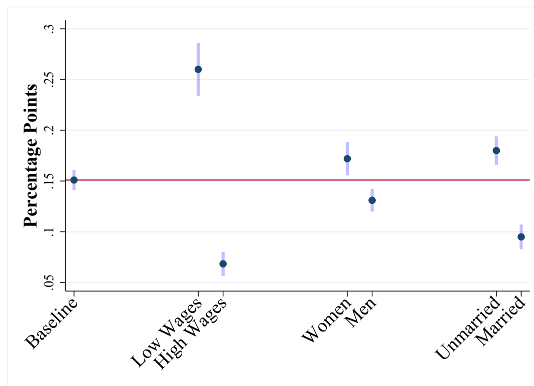
(b) Duration of Uninsurance



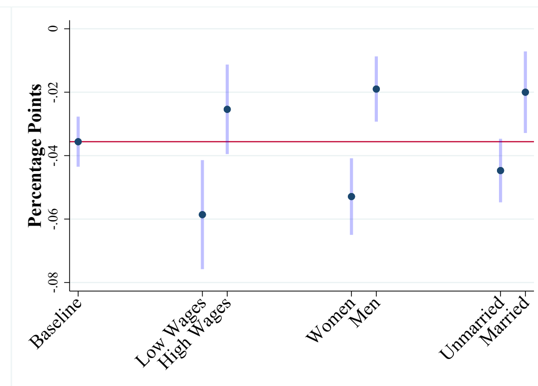
(c) Instantaneous Likelihood of Finding Coverage



(d) Employer Plan



(e) Medicaid Plan



(f) Marketplace Plan

Notes: This figure depicts estimates of the effect of Medicaid on different sub-populations as measured in the year prior to separation: the bottom and top 25th percentile of workers based on earnings (low and high wages), women and men, and individuals who were unmarried or married. Panel (a) reports estimates and 95% confidence intervals for the outcome measuring the likelihood of finding coverage one month after plan separation, panel (b) reports estimates for the duration of uninsurance, measured in months, and panel (c) reports estimates of the hazard ratio, or the instantaneous likelihood of finding coverage after plan separation, and panels (d)–(f) report estimates of the likelihood of an employer plan, Medicaid, or a marketplace plan serving as the first source of coverage. Point estimates are reported in Tables 7 and A7.

A. APPENDIX: FOR ONLINE PUBLICATION

Table A1
Monthly Source of Coverage by Medicaid Expansion, Unemployed: Transitions to $m+1$

	Expansion				Non-Expansion			
m	Uncovered	Employer	Marketplace	Medicaid	Uncovered	Employer	Marketplace	Medicaid
Uncovered	87%	9%	1%	3%	91%	7%	1%	1%
Employer	3%	96%	<0.5%	1%	4%	96%	<0.5%	<0.5 %
Marketplace	3%	5%	91%	2%	3%	5%	92%	<0.5%
Medicaid	2%	2%	<0.5%	96%	4%	2%	<0.5%	94%

Notes: This table describe the Markov Transition Matrix for policyholders who separate from an employer plan in 2016 and who are unemployed. Monthly coverage is observed for 24 months after separation. These statistics reflect the likelihood of transitioning across coverage sources from month m to month $m+1$.

Table A2
U.S. and Virginia, American Community Survey 2018

	U.S.	Virginia
Female (%)	50.8	50.8
Median Age	38.2	38.3
Race (%)		
White	75.1	71
Black or African American	14.1	21.2
American Indian or Alaska Native	1.7	1.1
Asian	6.8	8.1
Native Hawaiian and Other Pacific Islander	0.4	0.2
Other	5.5	3
Median Household Income ($)	61,937	72,577
Uninsured (% civilian, non-institutionalized)	8.9	8.8
Bachelor's or above (%)	32.6	39.3
Population	327,167,439	8,517,685

Notes: 2018 1-year estimates from American Community Survey. Data downloaded from https://data.census.gov/.

B. Full Population Results

For completeness, we provide all of our results using the full sample of policyholders that separate from an employer plan. Overall, the results follow similar patterns to what was found using the restricted sample of those who likely separated due to unexpected job loss.

Table A3 describes policyholders based on their duration of uninsurance following their policy separation. Each column reports statistics for policyholders by the number of months before new coverage is observed. We describe those who remain without new coverage for at least 24 months as "Never Covered."

Similar to the unemployed population, policyholders are less likely to be female and less likely to be joint filers as the duration of uninsurance increases. Prior wages decline with the duration of the spell, suggesting positive selection out of uninsurance, though 69% of those who remain uninsured two years later are employed in the following year. The share of people claiming UI benefits has a U-shape in duration of first coverage, suggesting that many of the transitions early on are by people choosing to switch jobs.

Table A4 describes the full population of separated policyholders based on the source of their new coverage. Employer health insurance serves a a first source for a larger share than in the unemployed sample: 77% vs 68%. This difference suggests that many plan separations may be due to job switches. Medicaid provides a first source of coverage roughly 10% of the sample, and the duration of uninsurance for those who purchase a Marketplace plan on average experience roughly one extra month of uninsurance.

Figure A1 shows the point-in-time composition of insurance for each month across the full 24 month post-separation window for all policyholders (panel a), policyholders living in an expansion state (panel b), and policyholders living in a non-expansion state (panel c) In each month, individuals are in one of four coverage groups: employer health plan, Medicaid, marketplace plan, or uncovered. These figures largely reflect the same dynamics seen for the unemployed population: Medicaid plays a large role in reducing uninsurance in expansion states.

Table A5 reports these estimates of the likelihood of regaining coverage one month after separation, the duration of uninsurance in months, and the Cox Proportional Hazard model estimates, respectively. The final three columns show the probability of different sources serving as the first new coverage.

Table A6 reports the six-month transitions across coverage types for Medicaid expansion and non-expansions states, respectively.The persistence of uninsurance is 5 percentage points higher in non-expansion states compared to expansion states (52% vs. 47%). Medicaid coverage appears to be more stable in expansion states, with 62% of those with Medicaid coverage in month m having Medicaid coverage 6 months later, compared with just 58% in non-expansion states.

Table A3
Summary Statistics by First Month of Coverage

	First Month of Coverage					
	1	3	6	12	24	Uncovered
Female	0.469	0.462	0.429	0.410	0.371	0.342
Married	0.477	0.352	0.296	0.290	0.254	0.274
Age						
18-25	0.0940	0.0781	0.0964	0.0985	0.109	0.0949
26-44	0.530	0.607	0.577	0.563	0.563	0.558
45-62	0.376	0.314	0.326	0.339	0.328	0.347
2015 Wages	66667.9	54233.2	50427.1	51482.5	46738.1	49155.2
Employed, 2017	0.927	0.939	0.911	0.905	0.805	0.687
Unemployed	0.122	0.179	0.276	0.258	0.232	0.187
2015 MAGI						
$<$ 100% FPL	0.0517	0.0326	0.0471	0.0515	0.0611	0.0687
100-138% FPL	0.0471	0.0480	0.0620	0.0646	0.0754	0.0779
138-400% FPL	0.424	0.553	0.575	0.567	0.597	0.583
$>$ 400% FPL	0.477	0.366	0.316	0.317	0.266	0.270
First Source of Coverage						
Employer Plan	0.835	0.811	0.751	0.762	0.731	.
Marketplace Plan	0.0278	0.0872	0.0869	0.0828	0.0777	.
Medicaid Plan	0.119	0.0740	0.138	0.125	0.136	.
N	7,036,614	589,964	160,753	69,938	21,080	551,314

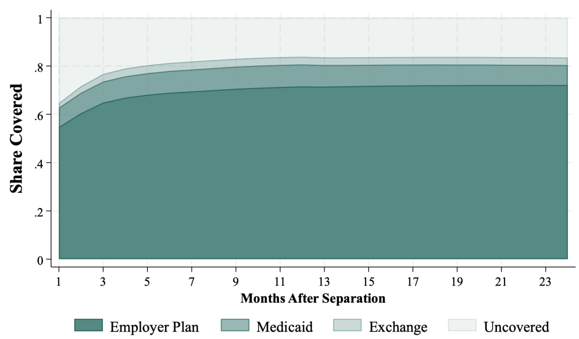
Notes: This table summarizes policyholders in 2016 of 2017 based on their first source of coverage after a separation from an employer plan in 2016. See also Table 1 notes.

Table A4
Source of New Coverage

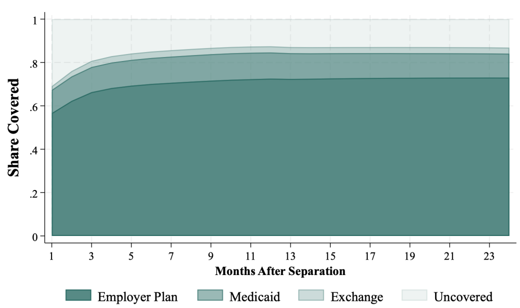
	First Source of Coverage		
	Employer	Medicaid	Marketplace
Female	0.452	0.575	0.515
Married	0.462	0.226	0.406
Age			
18-25	0.0881	0.150	0.0536
26-44	0.547	0.584	0.492
45-62	0.365	0.266	0.454
2015 Wages	67434.9	31254.5	52508.7
Employed, 2017	0.954	0.823	0.767
Unemployed	0.123	0.301	0.262
2015 MAGI			
$<$100% FPL	0.0290	0.202	0.0476
100-138% FPL	0.0364	0.156	0.0568
138-400% FPL	0.448	0.538	0.553
$<$400% FPL	0.487	0.103	0.343
Months of Uninsurance	2.479	2.754	3.940
N	8,269,789	1,168,444	475,348

Notes: This table summarizes policyholders in 2016 of 2017 based on their first source of coverage after a separation from an employer plan in 2016. See also Table 1 notes.

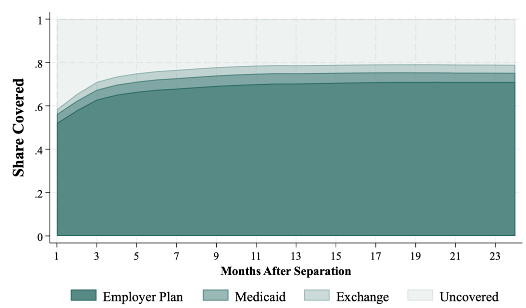
Figure A1. Post-Separation Coverage: All Policyholders



(a) All Policyholders



(b) Medicaid Expansion



(c) Medicaid Non-Expansion

Notes These figures depict the composition of sources of health insurance coverage in each of the 24 months after a policyholder separates from an employer plan in 2016.

Table A5
Post-Separation Coverage Dynamics:
Policyholders Who Separate From an Employer Plan in 2016

	Coverage Dynamics			First Source of Coverage		
	Covered at t=1 (1)	Months of Uninsurance (2)	Hazard Ratio (3)	Employer Plan (4)	Medicaid Plan (5)	Marketplace Plan (6)
Female	0.0557***	-1.357***	1.174***	-0.00953***	0.0691***	0.00697***
	(0.000401)	(0.00534)	(0.000986)	(0.000454)	(0.000281)	(0.000112)
Married	0.149***	-2.099***	1.302***	0.150***	-0.0164***	0.00183***
	(0.000562)	(0.00722)	(0.00116)	(0.000909)	(0.000511)	(0.000127)
Female *x* Married	-0.0296***	1.061***	0.889***	0.0352***	-0.0665***	-0.00255***
	(0.000644)	(0.00792)	(0.00114)	(0.000860)	(0.000481)	(0.000184)
Age	0.000974***	0.00192***	1.000	0.00184***	-0.00187***	0.000414***
	(0.0000178)	(0.000225)	(0.0000292)	(0.0000289)	(0.0000167)	(0.00000466)
Wages	0.0000255*	-0.000332*	1.000***	0.0000532*	-0.0000313*	-0.00000359*
	(0.0000112)	(0.000139)	(0.000000285)	(0.0000228)	(0.0000132)	(0.00000149)
Expansion State	0.102***	-1.340***	1.185***	0.0454***	0.0611***	-0.00534***
	(0.000309)	(0.00403)	(0.000766)	(0.000369)	(0.000198)	(0.0000914)
Unemployed	-0.161***	1.384***	0.841***	-0.254***	0.0818***	0.0102***
	(0.000430)	(0.00564)	(0.000746)	(0.000483)	(0.000330)	(0.000143)
Unconditional Mean	0.661	3.709	–	0.548	0.0798	0.0191
Month Fixed Effects	✓	✓	✓	✓	✓	✓
N	10,643,393	10,643,393	10,643,393	10,097,925	10,097,925	10,097,925

Notes: This table reports results results from estimating a OLS model (cols 1,2, 4, 5, 6) and a Cox Proportional Hazard model (col 3). Analysis is based on policyholders who separate from an employer plan in 2016. See also Table 4 and Section 3 for more details.

Table A6
Monthly Source of Coverage by Medicaid Expansion: Transitions to $m+6$

	Expansion				Non-Expansion			
m	Uncovered	Employer	Marketplace	Medicaid	Uncovered	Employer	Marketplace	Medicaid
Uncovered	47%	44%	2%	7%	52%	43%	3%	3%
Employer	6%	91%	1%	2%	8%	90%	1%	1%
Marketplace	7%	39%	50%	5%	9%	38%	52%	2%
Medicaid	7%	30%	1%	62%	12%	29%	1%	58%

Notes: This table describe the coverage transitions for all policyholders who separate from an employer plan in 2016 by whether they lived in an expansion state (left panel) or a non-expansion state (right panel). Monthly coverage is observed for 24 months after separation. These statistics reflect the likelihood of transitioning across coverage sources from month m to month $m+6$.

Table A7
Coverage Dynamics: Heterogeneity by Gender and Marital Status

	Coverage Dynamics			First Source of Coverage		
	Covered at t=1	Months of Uninsurance	Hazard Ratio	Employer Plan	Medicaid Plan	Marketplace Plan
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A. Women						
2019 × VA	0.0855***	-0.623***	1.211***	-0.108***	0.172***	-0.0529***
	(0.00917)	(0.0633)	(0.0269)	(0.00979)	(0.00837)	(0.00616)
Pre-Treatment Mean	0.414	4.00	–	0.733	0.156	0.0983
N	166,755	166,755	166,755	112,125	112,125	112,125
Panel B. Men						
2019 × VA	0.0478***	-0.357***	1.127***	-0.0932***	0.131***	-0.0190***
	(0.00879)	(0.0610)	(0.0246)	(0.00881)	(0.00567)	(0.00525)
Pre-Treatment Mean	0.402	4.180	–	0.817	0.0436	0.0656
N	189,866	189,866	189,866	124,989	124,989	124,989
Panel C. Unmarried						
2019 × VA	0.0836***	-0.659***	1.245***	-0.126***	0.180***	-0.0447***
	(0.00758)	(0.0548)	(0.0245)	(0.00877)	(0.00714)	(0.00511)
Pre-Treatment Mean	0.334	4.58	–	0.731	0.138	0.0875
N	234,040	234,040	234,040	146,405	146,405	146,405
Panel D. Married						
2019 × VA	0.0317**	-0.158*	1.051	-0.0493***	0.0950***	-0.0200**
	(0.0115)	(0.0735)	(0.0268)	(0.00972)	(0.00622)	(0.00656)
Pre-Treatment Mean	0.545	3.181	–	0.847	0.0362	0.0719
N	122,581	122,581	122,581	90,709	90,709	90,709

Notes: This table reports the effect of Medicaid Expansion on post-separation coverage dynamics using a difference-in-differences identification strategy. Panels A and B report estimates based on the subsample of female or male policyholders, respectively. Panels C and D report estimates based on the subsample of unmarried and married filers (as proxied by non-joint and joint filing status), respectively. Pre-treatment mean measured in Virginia in 2018. See Section 5 for more details.

C. Survival Models

If M is a non-negative random variable denoting time to regaining insurance, then its survivor function $S(m)$ is defined as follows

$$S(m) = 1 - F(m) = P(M > m),$$

and characterizes the probability of remaining uninsured after month m.

An empirical counterpart to the survivor function is a hazard function, $h(m)$, or the conditional failure rate. The hazard function describing the instantaneous likelihood of regaining coverage, conditional on an individual having been uninsured until month m can be written as follows:

$$h(m) = \lim_{\Delta m \to 0} \frac{Pr(m + \Delta m > M > m | M > m)}{\Delta m} = \frac{f(m)}{S(m)}.$$

We estimate $h(m)$ assuming a Cox proportional hazard model, which is a semiparametric model that is agnostic about the shape of the hazard function and assumes that covariates multiplicatively shift the baseline hazard function.

In discrete time, it is common that subjects are not observed from the onset of risk, $m = 0$. Indeed, this is the case in our dataset–we cannot observe periods of uninsurance that are smaller than one month given the discrete nature of our data. In other words, individuals who go uncovered for a matter of weeks between policies will appear to the econometrician as having regained coverage one month later. However, this does not affect the hazard function, which is an instantaneous rate that is not a function of the past.

The hazard rate for the jth individual in the data is

$$h(m|x_j) = h_0(m)\exp(x_j\beta_x)$$

We include the same covariates in the model for x_j as in the OLS model.

Finally, the interpretation of the estimated coefficients comes from the ratio of two individual hazards:

$$\frac{h(m|x_j)}{h(m|x_m)} = \frac{\exp(x_j\beta_x)}{\exp(x_m\beta_x)}.$$

Exponentiated coefficients are interpreted as the ratio of hazards for a one-unit change in the corresponding covariate. For example, the coefficient for a gender dummy variable, $female$, is interpreted as the ratio of the hazard for women compared to men. When $\hat{\beta} > 1$ ($\hat{\beta} < 1$), this implies that women are more (less) likely than men to regain coverage. Statistical significance is interpreted based on a null hypothesis that the exponentiated coefficient is equal to one. A rejection of this null hypothesis for the gender dummy would then imply that there is enough statistical evidence to reject the null hypothesis that women and men are equally likely to instantaneously regain coverage.