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ABSTRACT

Government Regulation and Lifecycle Wages: Evidence from Continuing Coverage Mandates*

We examine the lifecycle wage effects of health insurance market regulation that compels private insurers to offer continuing coverage to beneficiaries. Using a panel of male workers drawn from the National Longitudinal Survey of Youth 1979, we model wages across the lifecycle as a function of the mandated number of months of continuing coverage at labor market entrance. Access to continuing coverage is plausibly valuable to young workers as this benefit facilities job mobility, which is important for early career wage growth and lifecycle wages, but is costly to firms. We show that more generous mandated continuing coverage at labor market entrance causes an initial wage decline of roughly 1% that reverses after five years in the labor market leading to higher wages later in the career. Wage increases are observable up to 30 years after labor market entrance. We provide suggestive evidence that increased job mobility early in the career is a mechanism for the observed wage effects.

JEL Classification: J3, H2, I13

Keywords: regulation, job lock, continuing coverage, wage determination, persistence

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1. Introduction

Labor market entrance is a critically important period for establishing lifecycle wage profiles (Topel and Ward 1992, von Wachter 2010, Gardecki and Neumark 1998). During this period, workers gain skills and transition to higher paying jobs and/or jobs that offer a better worker-firm-specific match, which sets the worker's wage course for their career. Features of the labor market that stifle mobility can persistently lower wages. For example, a series of studies document that workers who enter the labor market during a period of reduced aggregate demand have persistently, but not permanently, lower wage and non-wage compensation (Schwandt and Von Wachter 2019, Kahn 2010, Oreopoulos, von Wachter, and Heisz 2012, Kondo 2015, Genda, Kondo, and Ohta 2010, Maclean 2014, Altonji, Kahn, and Speer 2016).[1]

We add to the literature by examining the importance of private health insurance market regulation for young male workers who are beginning their careers. In particular, we study the persistent effects of state-level government regulations that compel private insurers to include continuing coverage in insurance plans on wages across the lifecycle. Following regulation adoption, many firms that offer insurance to workers must therefore include the benefit, which can increase labor costs. Standard economic models of mandated benefits imply that, if workers value the wages, firms will reduce wages to offset benefit costs. The continuing coverage mandates we study allow separating workers and their dependents to maintain access to employer-sponsored health insurance (ESI) for a period of up to 20 months, with an average of 1.1 months, after leaving a job. These mandates, passed in the 1970s and 1980s, reflect the first regulation by state or Federal governments of continuing coverage in the U.S.

Continuing coverage is plausibly of particular value to young workers as it facilitates job

[1] We note that there is some heterogeneity across workers in wage effects. For instance, Hershbein (2012) documents at most modest, short term wage declines for lower-skill men. In general, wage effects appear to be strongest for high skill workers, in particular men.

mobility, which in turn promotes skill accumulation and improved worker-job matching, and thus higher wages across a worker's life. Additionally, because younger workers are more likely to change jobs than older workers (Groot and Verberne 1997), younger workers are potentially a demographically identifiable group to firms. For example, in our data, described later in the manuscript, we find that the propensity to change jobs in the past year is 55% for workers who have been in the labor market no more than five years and only 33% for those workers with more than five more years of experience. The ability to identify workers likely to value the benefit allows firms to pass costs onto this group through lower wages (Lahey 2012).

Given this background, we expect an initial negative relationship between mandated continuing coverage and wages, which will then become positive after workers are able to benefit from increased job mobility. Early career wages should be reduced to offset benefit cost. However, continuing coverage mandates are expected to increase wages as young workers spend time in the labor market and benefit from increased job mobility and ensuing improved matches.

We draw a long panel of workers from the National Longitudinal Survey of Youth 1979 (NLSY79). The NLSY79 is ideal for our research question as the dataset tracks workers from labor market entrance through mid-career. We estimate differences-in-differences (DD) style models with the NLSY79 to study continuing coverage mandate effects. The time period in which our sample entered the labor market, 1973 to 1991, coincides with the initial roll-out of state continuing coverage mandates. During this period 26 states implemented a continuing coverage mandate, offering us substantial policy variation that we can use for identification of wage effects. Another benefit of the NLSY79 is that it allows us to examine a time period prior to the large-scale movement towards self-insurance in the U.S. Self-insured firms are insulated from state-level insurance regulations through the Employee Retirement Income Security Act (ERISA) of 1974. In the second half of the time period in which our sample enters the labor market just 8% of firms self-insured (McDonnell et al. 1986).

We document an initial decline in wages for new labor market entrants attributable to mandated continuing coverage: an additional month of continuing coverage reduces wages by approximately 1% in the first five years in the labor market. Wages increase with experience for workers subject to these mandates at labor market entrance. After ten years in the labor market wages are 0.6% higher with each additional month of mandated continuing coverage at entrance and this effect persists through mid-career before dissipating after 30 years. We provide suggestive evidence that increased job mobility – proxied by the propensity to change jobs and cumulative number of jobs held – is a potential mechanism for the observed wage effects. The mobility finding mirrors reductions in job lock documented in previous work using samples of the working age males and exploiting the same source of variation that we leverage in our study (Gruber and Madrian 1994). However, we are the first study to show that policies minimizing job lock during important career development stages have lasting effects.

The paper proceeds as follows. A discussion of job lock and continuing coverage mandates is provided in Section 2. Our conceptual framework and hypotheses are presented in Section 3. Section 4 outlines our data, variables, and methods. Our main results and robustness checks are reported in Section 5. Section 6 concludes.

2. Related literature

The U.S. health insurance system is characterized by a patchwork of targeted public programs (e.g., Medicare for the elderly and Medicaid for the poor), ESI, and insurance purchased by individuals. ESI is historically the most common form of insurance for non-elderly adults; 77% of insured non-elderly adults held ESI coverage in 2009, the year before implementation of the Affordable Care Act (ACA), which led to largescale increases in both public and private coverage.[2] Private individually purchased insurance is generally more costly and less generous in terms of covered benefits and cost-sharing than ESI; see, for example, Table 2 in Gruber and Madrian (1994) which covers our study period. The quality of public insurance available to the non-elderly (e.g., Medicaid) is a perennial concern among policymakers (Decker 2012) and many low-income individuals were not eligible for public insurance during our study period due to categorical requirements (Buchmueller, Ham, and Shore-Sheppard 2015). A non-trivial share of the U.S. population, particularly prior to the ACA, is uninsured as many low-income individuals are not eligible for public coverage but cannot afford private coverage or, for myriad reasons, decide not to purchase/enroll in insurance.

Heavy reliance on employers for insurance coverage leads to concerns of 'job lock' among policy makers; that is workers remain in undesirable jobs to retain health insurance coverage, causing reduced job mobility and worker-firm match quality. A robust economic literature has examined the existence of job lock and factors that mitigate or exacerbate this phenomenon. Comprehensively reviewing this large and historic literature is beyond the scope of our study, instead we simply summarize studies closely related to our own work.

Perhaps the most relevant work for our study is Gruber and Madrian (1994), which investigates the relationship between years of mandated continuing coverage and job turnover among prime-age males. The authors estimate an increase in turnover of 9% resulting from an additional year of mandated continuing coverage, implying a strong relationship between job lock and the source of variation we examine in this manuscript. In particular, the authors calculate that a year of continuing coverage reduces job lock by 40%.

There is a much broader literature on the general relationship between health insurance and labor supply/mobility (e.g., own employment transitions, spousal transitions, retirement)

[2] Authors' calculation among insured non-elderly adults using the 2009 American Community Survey (Ruggles et al. 2017). Details available on request from the corresponding author.

that relates to our work. Many studies document evidence of job lock (Gruber and Madrian 1997, Rashad and Sarpong 2008, Hamersma and Kim 2009, Gruber and Madrian 2002, Garthwaite, Gross, and Notowidigdo 2014, Kofoed and Frasier 2019, Dave et al. 2015, Baicker et al. 2014, Boyle and Lahey 2010, Gruber and Madrian 1995, Madrian 1994), although there are exceptions (Bailey and Chorniy 2016, Heim, Lurie, and Simon 2018, Kaestner et al. 2017).

Heterogeneity in estimated job lock effects is not surprising as different studies rely on different sources of variation that may affect different types of workers (e.g., increases in Medicaid coverage vs. changes in private insurance) and study different time periods experiencing different economic conditions which may influence available jobs. Our contribution to this literature is to study both the contemporaneous and lifecycle effects among new labor market entrants. This question, to the best of our knowledge, is unstudied.

3. Conceptual framework and hypotheses

Summers (1989) provides one of the early economic analyses of health insurance mandates on wages and employment; we focus mainly on wage predictions from the Summers framework as that is most salient to our study.[3] Pre-mandate, the labor market is in equilibrium at the intersection D_1 and S_1, with employment level E_1 and wages of W_1 as depicted in Figure 1. Summers, assuming that the cost of the mandate is a per hour rate, argues that health insurance mandates increase labor costs and should, all else equal, lead to a decrease in demand for labor among firms by the cost of the mandate. This reduction in labor demand is depicted as a shift from D_1 to D_2. Thus, the mandate should lead to a lower level of employment and wages (E_2',W_2'). If workers value the mandate, however, then this valuation will lead to an increase in labor supply as workers enter the market to gain access to the benefit (from S_1 to S_2). The labor

[3] The Summers model can explain a wider range of mandated benefits, but this framework is commonly employed within the health insurance mandates literature and we follow that tradition for our discussion.

supply increase will have two effects: attenuate the employment drop and increase the wage decline. At the new equilibrium, wages and employment will fall (E_2,W_2). The extent to which the mandate affects wages and employment is determined by workers' valuation of the benefit.

If workers fully value the benefit, the incidence of the mandate will be entirely passed on to the workers in terms of lower wages and there will be no impact on employment levels (as firms experience no increase in labor costs). Alternatively, if workers do not value the benefit to any extent, the cost of the mandate will be fully born by firms, wages will be unchanged, and overall employment will decline to fully offset the mandate cost. Intermediate valuations of the benefit by workers will lead to both lower wages and employment levels, with the relative magnitudes of these effects determined by worker preferences. Summers (1989) notes that features of the U.S. labor market, such as minimum wages and anti-discrimination laws, may limit firms' ability to reduce wages to offset mandate costs, which complicates predictions.

The Summers model offers important insight into the wage effects of mandated benefits. However, the model is static and therefore does not offer direct predictions into persistent effects of continuing coverage mandates for new labor market entrants, which is important for our study. We next use intuition from the job lock literature to formulate a simple model of lifecycle wages in the presence of mandated continuing coverage at labor market entrance.[4]

Assuming a competitive labor market, in a job that does not offer fringe benefits, worker *i* earns wage equal to their marginal product: $w_i = MP_i$. If the firm elects to offer (actuarially fair) health insurance and can accurately adjust overall compensation for worker *i* by his expected healthcare expenditures ($E[HC_i(M)] > 0$), where M is the number of months of continuing coverage mandated by state law, then the worker would receive $\dot{w}_i = MP_i -$

[4] To develop this model, we also borrow logic from a static wage determination framework proposed by Bhattacharya and Bundorf (2009) in the context of obesity and healthcare costs.

$E[HC_i(M)]$ if the firm fully passes on insurance costs to the worker. As noted above, the pass-through rate is determined by worker valuation of the mandated benefit (Summers 1989). Gruber (2000) argues that firms cannot accurately risk-adjust premiums and wages at the worker-level, and instead reduce each worker's wage by the average expected healthcare costs among all J workers at the firm: $\frac{1}{J} * \sum_{j=1}^{J} E[HC_j(M)] = \overline{HC(M)} > 0$. This behavior implies that each worker receives insurance and the following wage: $\dot{w}_i = MP_i - \overline{HC(M)}$. Because continuing coverage should facilitate job mobility and wage growth, we expect that wage effects will vary with time in the labor market (e): $\ddot{w}_{i,e} = MP_{i,e} - \overline{HC(M)} + B_{i,e}(M)$. Where $B_{i,e}(M)$ is the change in wages from better job matches attributable to mandated continuing coverage at labor market entrance and is assumed to be non-negative and likely positive over time. We note that $\overline{HC(M)}$ may also change with time in the labor market if, after a job transition, the expected average healthcare costs at the firm are different.

Based on this simple dynamic framework, we hypothesize that months of continuing coverage at labor market entrance will:

H1: Reduce wages initially.

H2: Increase wages later in the lifecycle.

H3: Increase job mobility early in the lifecycle.

We will test these hypotheses in a sample of workers drawn from the NLSY79.

4. Data, variables, and methods

4.1 Data

We draw data on a panel of workers from the NLSY79. We retain only men in our main analyses because we focus on an older cohort, and we prefer a sample with high labor force attachment for which the mandates we study may have some 'bite.' The original NLSY79 sample consists of 12,686 youth ages 14 to 22 in 1979. The survey was administered annually between 1979 and 1993, and bi-annually from 1994 onward. These data are well suited to our research question as the NLSY79 was designed to track a cohort of workers as they transitioned into the labor market and throughout their career. We are able to follow workers from labor market entrance, which we define as the year after leaving school, through mid-career. We include all types of school-leaving: dropouts and those completing diplomas/degrees. We truncate the study period in 2012, when workers are in their early-to mid-50s, to avoid confounding from the ACA, which fundamentally altered the healthcare delivery system in the U.S. and, in particular, the tight link between insurance and employment among non-elderly adults (Oberlander 2010). Core provisions of this Federal Act went into effect in January 2014.

We consider the hourly wage which is available in all years; we inflate nominal wages to 2012 dollars using the Consumer Price Index (CPI). We exclude workers with real wages less than $1 per hour and those with real wages greater than $1,000 per hour. We take the logarithm of wages, thus regression coefficient estimates have the interpretation of an approximation to the percent change. To test our hypothesis that continuing coverage promotes wage growth by reducing job lock and increasing labor mobility, we also consider job changes and the cumulative number of jobs held by a worker in each year of the survey. We do not discriminate between voluntary and involuntary job changes.

We include men who enter the labor market between the ages of 16 and 30 years, meaning that our analysis sample includes men between 16 and 54 years. Compulsory schooling laws prevented legal school leaving before age 16 in most cases during the period in which our sample entered the labor market and there are very few individuals who left school after age 30 in the NLSY79. These exclusions, and others that are detailed later in the manuscript, lead us to an analysis sample that includes 5,161 unique workers – or 81% of the full male sample – and 76,694 worker/year pairs.

4.2 Linking mandated months of continuing coverage at labor market entrance to the NLSY79

We use the state and year in which each individual leaves school as a measure of the location and period they enter the labor market. A limitation of the NLSY79 data is that we only have state of residence beginning in 1979, but many individuals in our sample enter the labor market before this year. Thus, we use state of residence at age 14 as a proxy for the labor market entrance state. Workers who reside outside the U.S. at age 14 are excluded. If state of residence at age 14 is missing, we impute this state with birth state if the respondent was born in the U.S. We note that using the state of residence at age 14 leads to measurement error for those individuals who migrate across state lines between age 14 and labor market entrance. However, as we report later in the manuscript, our results are insensitive to alternative imputation approaches. We refer to state of residence at age 14 as the 'labor market entrance' state.

We define the year of labor market entrance using retrospective information on school-leaving collected between 1979 and 1998 (Maclean 2013). Non-enrolled NLSY79 respondents are asked to report the year in which they left school. If a respondent indicates that they completed no formal education, we exclude them from the analysis sample. We focus on the sample of workers who enter the labor market between 1973 and 1991.[5] Following Gruber (1994b) and Gruber and Madrian (1994), we exclude respondents who enter the labor market in DC, Hawaii, and West Virginia. Table 1 reports the number of respondents entering the labor market by year. The Federal Consolidated Omnibus Budget Reconciliation Act of 1985 (COBRA), which supersedes less generous state laws, mandates 18 months of continuing coverage for most employees. This Act was fully effective in July 1987 (Gruber and Madrian 1994). 97% of our sample enters the labor market before this Federal Act was in place and thus

[5] Starting the study period in 1973 is an artifact of our sample exclusion of men who left school before age 16. Details available on request from the corresponding author.

COBRA is unlikely to influence our findings.[6] However, we address the empirical importance of COBRA in robustness checking later in the manuscript.

Table 2 reports the effective date for each continuing coverage mandate using legal data collected by Gruber and Madrian (1994). We also list the number of mandated months of coverage. Some mandates have mid-year effective dates. For such mandates, we code the law as effective in year *t* if it became effective between January and June, and effective in year *t+1* if the effective date was July to December. The majority of respondents in our sample entered the labor market in the first half of the year (74%, details available on request from the corresponding author), leading us to this coding scheme. Figure 2 reports the count of new or amended state-level continuing coverage state mandates in each year or our study.

Minnesota was the first state to implement a continuing coverage mandate, the state implemented legislation compelling private insurers to cover six months of coverage in 1974. This mandate was strengthened to 12 months of coverage in 1983 and then 18 months in 1987. Other early adopting states were Connecticut (1975) and Oklahoma (1976). The mandated months of coverage ranges from one to 20 months, with a conditional mean of 7.6 months.

4.2 Controls

We include a set of pre-determined individual-level variables in our regression models: race/ethnicity, age at labor market entrance, level of education at labor market entrance, ability (age-standardized Armed Forces Qualification Test [AFQT]), and parental education (mother's and father's years of education). We include indicators for missing covariates and assign missing observations to the sample mean (continuous variable) or mode (binary variable). Results, reported later, are not appreciably different if we exclude these controls.

[6] COBRA was phased in between July 1986 and June 1987 (Gruber and Madrian 1994). Most individuals graduate from school, which we define as labor market entrance, in the first six months of the year (74% in our sample) and are therefore not likely affected by changes that occurred in 1986.

We also include state demographic information on sex and age from the Survey of Epidemiology and End Results (SEER) made available through the National Bureau of Economic Research. State political preferences are important predictors of health policies (Courtemanche et al. 2017, Maclean et al. 2018). To proxy political preferences, we use a measure of state citizen ideology developed by Berry et al. (1998).[7] Broadly, for each state this index reflects the ideological ranking of each member of Congress and each district. Lower scores indicate more conservative ideology within the state. We also include the employment-to-population ratio, per capital income (inflated to 2012 dollars using the CPI), and state population from the Bureau of Economic Activity. Finally, we include the count of 'high cost' private insurance mandates as defined by Gruber (1994b): alcohol use disorder treatment, illicit drug use disorder treatment, mental illness treatment, and chiropractor services.[8] A concern may be that the continuing coverage mandates we study influence the employment level (e.g., mandates reduce overall labor demand). However, as we document later in the manuscript, our results are robust to excluding labor market entrance state-level controls.

4.3 Empirical model

Equation (1) presents the DD-style regression model we use to estimate the effect of months of mandated continuing coverage at labor market entrance on wages across the lifecycle:

(1) $W_{i,s,t,g} = \alpha_0 + \sum_{j=1}^{7} \beta_j M_{s,t} + X_i \alpha_1 + C_{s,t} \alpha_2 + \theta_g + \gamma_s + \tau_t + \varepsilon_{i,s,t,g}$

$W_{i,s,t,g}$ is the logarithm of the hourly wage measured for individual *i* in labor market entrance state *s* and labor market entrance year *t* measured in survey year *g*. $M_{s,t}$ is the number of months of continuing coverage (lagged one year) in labor market entrance state *s* and labor market entrance year *t*. We interact this variable with indicators for *j* years of labor market

[7] We use an updated version of this coding system made available by Robert Fording. Please see https://rcfording.wordpress.com/state-ideology-data/; accessed February 20th, 2019.
[8] We note that Gruber classifies continuing coverage as a high cost mandate as well.

experience where j is 1 to 5, 6 to 10, 11 to 15, 16 to 20, 21 to 25, 26 to 30, and 31 or more years.[9] Lagging the mandate one year allows time for insurance contracts, which are typically renewed annually, to incorporate new regulations (Maclean, Popovici, and Stern 2018).

Labor market experience is proxied with the difference between the labor market entrance year and the year in which the outcome variable is measured (survey year) (Kahn 2010, Maclean 2013). X_i is a vector of personal characteristics and $C_{s,t}$ is a vector of labor market entrance state characteristics. θ_g includes survey year fixed-effects. γ_s is a vector of labor market entrance state fixed-effects and τ_t is a vector of labor market entrance year fixed-effects. Labor market entrance state fixed-effects account for time-invariant state characteristics that influence both the propensity for a state to adopt continuing coverage mandates and wages of new labor market entrants while labor market entrance year fixed-effects control for secular trends in wages that affect the nation as a whole.

We apply weighted least squares using NLSY79 sample weights, although as we document in robustness checking are our results are not sensitive to removing the weights. Standard errors are clustered around the labor market entrance state; we have 48 states which is sufficient for consistent estimation of standard errors (Cameron and Miller 2015).

We leverage variation in the number of months of continuing coverage within states over time to identify wage effects. The key identifying assumption of our DD model is:

$$(2) \qquad Cov\left(M_{s,t}, \varepsilon_{i,s,t,g} \middle| X_i, C_{s,t}, \theta_g, \gamma_s, \tau_t\right) = 0$$

In words, after conditioning on variables included in Equation (1), continuing coverage mandates are as good as randomly assigned. We examine the empirical importance of common treats to identification later in the manuscript.

5. Results

[9] We have estimated more parametric versions of this specification and results, available on request, are similar.

5.1 Summary statistics

Summary statistics are reported in Table 3. The average hourly wage in our sample is $22.38. Men in our sample enter labor markets in which private insurers were compelled offer just over one month of continuing coverage. On average, men are just under 19 years of age at labor market entrance and the majority has completed less than a college degree. The demographics of the sample are comparable to an older sample such as the NLSY79. For instance, the sample is less racially and ethnically diverse than the current U.S. population.

5.2 Validity and establishing contemporaneous effects of continuing coverage mandates

We estimate DD-style models to study the persistent effects of high cost mandate on wages across the lifecycle. The key assumption required for these models to recover estimates of causal effects is the 'parallel trends' assumption. We must assume that treatment states (i.e., those than passed months of continuing coverage) and comparison states (i.e., those that did not pass months of continuing coverage) would have followed similar trends in wages had mandates not been passed. This assumption is untestable but it is standard in the literature to provide suggestive evidence by examining trends in outcomes prior to mandate passage.

We estimate an event study model in the NLSY79 in the spirit of Autor (2003). We first center the NLSY79 data around the year in which a state implemented a continuing coverage mandate for the first time (see Table 2).[10] We then form a series of single-year 'leads' and 'lags' around the first continuing coverage mandate from seven years in advance of the mandate to twelve years after the mandate. Cell sizes, in event time, more than twelve years following the initial mandate become small and we exclude more distal post-event years (Lovenheim 2009). We code non-adopting states as zero for all mandate leads and lags. The omitted period is one year prior to mandate adoption. Our event study is outlined in Equation (2):

[10] We impose fewer restrictions on the sample (e.g., we do not require a valid labor market entrance year and state for inclusions) to maximize sample size. Details available on request from the corresponding author.

(2) $W_{i,j,g} = \partial_0 + \sum_{j=-7}^{-2} \delta_j Lead_j + \sum_{k=0}^{12} \delta_k lag_k + X_i \partial_1 + C_{j,g} \partial_2 + \varphi_j + \theta_t + \mu_{i,j,g}$

All other variables are as defined in Equation (1), however, we include current state (j) fixed-effects (φ_j) rather than labor market entrance state fixed-effects as we focus on contemporaneous effects in the event study. Results are reported graphically in Figure 3.

We do not observe statistically significant evidence of differential pre-trends between the adopting and non-adopting states; coefficient estimates on the lead variables are small in magnitude, statistically indistinguishable from zero, and change signs. Examination of the lag coefficient estimates suggests that wages decline in adopting states after private insurers are compelled to include continuing coverage as a benefit. We note that, because our sample declines in size and event studies are data hungry, some lag point estimates are imprecise.

We next turn to the Current Population Survey Annual Social and Economic Supplement (ASEC) to estimate the contemporaneous effects of continuing coverage mandates in a general sample of workers before proceeding to the NLSY79 in which we have one just cohort of workers that we track from labor market entrance through mid-career. The purpose of this exercise is to confirm that (i) the mandates we study affect wages contemporaneously and (ii) these effects are observable among younger workers (those likely to be new to the labor market). We use the ASEC as it is large-scale state representative dataset that is commonly employed in both the mandated benefit and job lock literatures (Kaestner and Simon 2002, Gruber 1994b, a, Boyle and Lahey 2010). We include adults ages 16 to 54 years over the roughly the same period in which our NLSY79 sample entered the labor market: 1975 to 1991 (wage data, which pertains to the past calendar year, for these years is available in the 1976 to 1992 files).

Reported wages are not available in all years of the ASEC, thus we construct hourly wage information using past year annual wages and salary earnings, weeks worked, and usual hours worked per week. We note that constructing wages in this manner likely leads to measurement error which is a limitation. Prior to 1975, we are not able to measure hourly wages in the ASEC with any degree of accuracy (details available on request). We view our ability to accurately measure hourly wages in all years in the NLSY79 as a further advantage of that dataset. We estimate an equation similar to that outlined in Equation (2) with the exception that estimate a DD-style model rather than an event study and we have fewer individual-level controls in the ASEC than the NLSY79. We lag months of continuing coverage by one year.

Results are reported in Table 4. We find that an additional month of mandated continuing coverage reduces wages by 0.2%. We stratify the sample by age: 16 to 30 years and 31 to 54 years. Stratifying the sample in this manner allows us to focus on workers more likely to be new labor market entrances and established workers. Mandate effects are only statistically different from zero in the younger worker sample. This finding offers premise for our hypothesis that younger workers, because they are more likely to change jobs to increase wages, value mandated continuing coverage and are potentially demographically identifiable to firms, allowing firms to pass costs associated with this mandated benefit to those workers. Further, younger workers may place more value on this benefit than older workers.

5.3 Lifecycle effects of mandated continuing coverage on wages

Table 5A reports estimates of the effect of months of continuing coverage at labor market entrance on hourly wages over the lifecycle. Men who enter labor markets with more generous continuing coverage initially earn lower wages, but this effects dissipates after six to ten years in the labor market and then leads to higher wages. In particular, an additional month of mandated continuing coverage reduces wages by 1.1% in the first five years in the labor market,[11] after that time wages increase by just under 1% in each five-year experience bin. Increased wages are observable from six to 30 years after labor market entrance.

[11] The 95% confidence of this point estimate overlaps with the 95% confidence interval surrounding the point estimate in the CPS regressions (section 5.2).

We hypothesize that mandated continuing coverage increases lifecycle wages primarily by reducing job lock early in the career and facilitating mobility between jobs among young workers. We attempt to shed light on the empirical importance of this pathway by examining the effect of continuing coverage benefits on the probability of a past year job change.[12] Results are reported in Table 5B. We observe that an additional month of mandated continuing coverage increases the probability of a job change by 0.8 percentage points or 2.2% in the first five years in the labor market, with effects dissipating and becoming statistically indistinguishable from zero after that level of experience. This pattern of results suggests that continuing coverage facilitates job changes by young workers which, in turn, promotes wage growth.

We also explore the effect of continuing coverage on the cumulative number of jobs held across the career (Table 5B). We observe that six to ten years after labor market entrance the number of cumulative jobs is 0.8% higher (0.06 units) for those workers who entered the labor market with an additional month of continuing coverage. No other coefficients are statistically different from zero, although all carry a positive sign. Our findings for cumulative jobs supports our job change findings: mandated continuing coverage increases job mobility, and hence the number of jobs held, and the effects are concentrated early in the career, precisely at the same time period when we observe a decline in wages.

5.4 Heterogeneity by worker skill and minority status

We next estimate separate regressions for workers of different skill levels (Table 6A) and minority groups (Table 6B). We focus on workers who enter the labor market with some college education (‘higher skill workers’) and a high school diploma or less (‘lower skill workers’). We

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[12] We calculate our job change variable using the count of jobs in adjacent years. For instance, to construct the variable for year *t*, we take the difference between the cumulative number of jobs reported in *t* and the cumulative number of jobs held in *t-1*. If the number of cumulative jobs is greater in *t* than the cumulative number of jobs in *t-1*, we classify the respondent as having changed jobs. Cumulative number of jobs is weakly monotonic across survey years. Because we require information on cumulative jobs in two adjacent years to construct this measure, we do not have a value of this variable for the first year of the survey (1979).

separate workers into two minority status groups: white (non-minority), and African American and Hispanic (minority). We do not have the sample size to separately analyze African Americans and Hispanics. Both higher and lower skill workers who enter labor markets with more generous continuing coverage mandates experience initial wage declines followed by wage increases, although the initial decline is only statistically distinguishable from zero for lower skill workers. Our sample size for higher skill workers is somewhat small and thus differences in precision may be attributable to power differences. Both white and minority workers experience an initial decline in wages attributable to more generous continuing coverage, however, only white workers experience a wage increase as they gain experience. For minority workers, wages decline initially when continuing coverage mandates are more generous and there is no corresponding benefit that occurs later in the career.

5.5 Lifecycle effects on additional labor market outcomes

We are most interested in estimating the effects of mandated continuing coverage on wages in this study. However, the mandated benefit literature shows that firms may adjust to increased labor costs associated with implementation of months of continuing coverage in other ways (Cutler and Madrian 1998, Kaestner and Simon 2002, Lahey 2012). Failure to consider these alternative margins of adjustment may lead to an inaccurate assessment of the overall lifecycle effects of months of continuing coverage for new labor market entrants. To investigate other margins we examine: ESI offers,[13] and the logarithm of annual weeks worked, usual hours worked per week, and annual earnings. If insurance becomes too costly after benefit coverage is mandated, some firms may cease offering ESI (Sloan and Conover 1998). On the other hand, as insurance becomes more costly, firms may extract extra more labor from each worker (Cutler

[13] Unlike many surveys used in the mandated benefit literature, for instance the ASEC, the NLSY79 includes information about an ESI offer, not whether the respondent holds ESI. This distinction is valuable as we are able, in the NLSY79, to avoid issues related to endogenous take-up of ESI which may be related to continuing coverage. For instance, if premiums rise following a mandated benefit, some workers may elect to drop ESI coverage.

and Madrian 1998). Examination of earnings allows us to assess the overall effect of changes in wages and labor supply on a worker's compensation. Results are listed in Table 7.

Across all outcomes we consider, there is an initial decline that dissipates after five years in the labor market. However, we do not observe corresponding increases later in the career for most outcomes we consider. An exception is annual earnings: the lifecycle pattern of effects mirror that of wages with an initial decline followed by increased earnings later in the career. The initial decline is somewhat larger than that observed for wages (2.9% one to five years after labor market entrance) but later effects are more comparable (roughly 1% to 2% increase). The larger initial decline is potentially explained by the reduction in both wages and labor supply, with the latter effect dissipating after five years in the labor market. However, examination of the estimated effects implied by the tails of our 95% confidence intervals surrounding our wages and earnings point estimates suggests more similar effect sizes, thus we do not wish to overstate any heterogeneity across earnings and wages. We note that there is a decrease (increase) in the probability of ESI (usual number of hours worked per week) later in the career. We are uncertain as to the mechanisms that lead to these changes.

5.6 Robustness checks

We estimate a number of checks on our specification and sample to assess the stability of our findings. Findings that are sensitive to reasonable changes to these study features raise concerns that any findings are spurious. Reassuringly, our results are broadly robust. For brevity, we note the specific checks and discuss important deviations from our main findings.

A primary concern with the NLSY79 is that the dataset is nationally representative is not representative at the level of our treatment variable (i.e., labor market entrance state) which can lead to inaccurate estimates (Maclean, Tello-Trillo, and Webber 2019). To address this concern, we re-estimate Equation (1) using the ASEC 1976 to 2013 (which covers wages over the period

1975 to 2012). The ASEC is representative at the level of the state and is thus not subject to the above noted concern. We make some changes in specification and variable definitions due to differences across datasets. In particular, we impute labor market entrance year following Genda, Kondo, and Ohta (2010), use a smaller set of individual-controls, and use current state as a proxy for labor market entrance state. Moreover, we construct a proxy for hourly wages as discussed in Section 5.2; while we have access to actual hourly wages for a subset of years of the data we chose to use the constructed wages to ensure a consistent outcome variable across all years. Full details are available on request from the corresponding author. Results, reported in Table 8A, are similar to our NLSY79 findings. We note that in the ASEC, and not the NLSY79, that wages appear to decline even after 30 years in the labor market. While we cannot fully assess the factors that lead to this finding, we hypotheses that workers, who are approaching standard retirement ages, may begin to phase into retirement by taking lower wage jobs. Their increased wages across the lifecycle may afford them this ability later in life. Alternatively, sample sizes in the NLSY79 decline as the cohort ages, thus – while the point estimate on the most distal mandate-bin interaction (31 or more years) carries a negative sign – it is imprecise. Within the NLSY79, we exclude observations with potentially poor coverage, in particular those observations with less than 20 respondents in a given labor market entrance state/year combination following Kondo (2015). Results are similar to our main findings (Table 8B).

A second concern is COBRA, a Federal Act that compelled private insurers to include 18 months of continuing coverage, that was implemented between July 1986 and June 1987. This Act rolled out as roughly 3% of our sample enter the labor market. We explore the empirical importance of this Act for our study in two ways. First, we exclude respondents who enter the labor market after 1987, when COBRA was fully effective and implemented across the nation, and second (retaining all observations) we code all states as requiring nine months of continuing coverage in 1986 and 18 months in 1987 and beyond (with the exception of Connecticut that requires 20 months in these years) to account for the Federal Act phase-in period. Results (reported in Table 9) are not sensitive to these changes. Given that the vast majority (97%) of our sample leaves school pre-COBRA, we interpret our findings as capturing the effects of state regulation of continuing coverage and not Federal COBRA.

We alter the control variables included in Equation (1) (Table 10A). In particular, we (i) exclude individual- and time-varying state-level controls, (ii) include labor market entrance division-by-year fixed-effects (we use the nine U.S. Census divisions), (iv) include labor market entrance state-specific linear trends, (v) do not lag the continuing coverage mandates one year (instead we use the contemporaneous mandates), and (vi) include survey state fixed-effects. We also explore the sensitivity of our findings to different weighting schemes and sample criteria (Table 10B). In particular, we estimate unweighted regressions and exclude those respondents who entered the labor market prior to 1974 (the year in which ERISA became effective). We also use different imputation approaches to assign labor market entrance states. We use (i) state of residence in 1979 (the first year of the NLSY79) and (ii) birth state (Table 10C).

We next sequentially exclude each state that adopted a continuing coverage mandate during our study period from the analysis sample and re-estimate Equation (1). This analysis explores the extent to which any particular state drives our findings (Tables 11A-11G). Results suggest that our main findings are not determined by unique experiences of one or more states. We regress the number of mandated continuing coverage months on time-varying state characteristics included in Equation (1), and state and year fixed-effects (Table 12). Pei et al (2019) note that such a test of balance offers useful insight on whether or not the conditional independence assumption holds. Overall, we find little evidence that these factors predict our treatment variable. However, we note that states with a larger share of older residents (higher per capita income) have more (fewer) months of mandated continuing coverage. This finding may suggest that states with older populations (and thus more likely to have access to Medicare) are less likely to support mandated continuing coverage mandates while higher income states are more likely to value this benefit through altruism or some other factor. Reassuring, our results are robust to excluding these controls.

Finally, we regress pre-determined and time-invariant characteristics of labor market entrants on the number of months of mandated continuing coverage at labor market entrance (white race, rural residence at age 14, born outside the U.S., and mother's education and father's education measured when the respondent was 14 years of age). This exercise allows us to test whether our treatment variable lead to compositional change among new labor market entrants which could confound our estimates (Table 14A). We use the first year of data for which we observe a respondent in the NLSY79, thus we have a smaller sample size. We find no evidence to support compositional change among labor market entrants attributable to the mandate. We also show that our results are robust to including these covariates in Equation (1); see Table 14B.

7. Discussion

We explore the lifecycle effects of state government regulation that compels private insurers to offer continuing coverage on new labor market entrants' wages. To the best of our knowledge, we are the first to study this question. Thus, we add to the literatures on economic shocks experienced at labor market entrance, job lock, and on wage offsets attributable to mandated benefits. Continuing coverage is plausibly particularly valuable to new labor market entrants as this regulation has been shown to alleviate job lock in the general population of workers (Gruber and Madrian 1994), and the ability to move from job-to-job promotes wage growth and better job-firm matches which will have persistent and positive effects on wage profiles based on theories of career development and empirical work. Further, firms may recognize that younger workers are more likely to change jobs than older workers, suggesting that younger workers are an identifiable group, which can allow firms to pass on costs to workers who may value the benefit.

Several findings emerge from our analysis. First, continuing coverage mandates at labor market entrance reduce wages: an additional month of continuing coverage reduces hourly wages by roughly 1%. The hourly wage rate in our sample is $22.38 in 2012 dollars and the mean number of months of mandated continuing coverage is 1.1 months over our study period. Thus, for the average worker these regulations lead to a 27 cent reduction in the hourly wage over the first five years in the labor market. Effect sizes are arguably reasonable given the high cost of health insurance coverage, the annual cost of an ESI family plan was $5,137 in 1987 (Gabel et al. 1988), with the costs of individually purchased insurance – the most plausibly outside insurance option for workers – was even higher.[14] Further, Handel (2013) estimates a valuation of more than $2,000 to continuing on the same health insurance plan even in the case when the worker remains with the same firm. Finally, we study a period in advance of large-scale self-insurance, just 8% of firms self-insured over our study period, and thus the mandates we study likely have more 'bite' than comparable state-level insurance regulations in today's labor markets. Collectively, given the high costs of health insurance, the scarcity of public options for health insurance at the time (Buchmueller, Ham, and Shore-Sheppard 2015), the value workers place on retaining their coverage, and the very limited self-insurance by firms in the 1970s and 1980s, our effect sizes seem reasonable for an intent-to-treat (ITT) estimate.

Second, over time workers who entered the labor market with more generous continuing coverage benefits earn more than their counterparts, which is in line with increased job mobility and, in turn, better job-firm matches. Third, we show that greater job mobility is an important

[14] Inflated by the authors from the original estimate of $2,520 in 1987 dollars to 2012 dollars using the CPI.

pathway from mandated continuing coverage to higher wages later in life: workers who enter the labor markets characterized by more generous continuing coverage are more likely to change jobs in early career and work in more jobs over their lifecycle. Fourth, there is heterogeneity in mandate effects across types of workers. Finally, mandate effects are not localized to wages, we observe some mandate effects on ESI offers, measures of labor supply, and annual earnings.

Our study has limitations. First, we focus on an older cohort of U.S. workers. Therefore, the generalizability of our findings to different cohorts and countries is not clear. However, in terms of relevance for the U.S., we note that continuing coverage is codified into insurance coverage by the ACA, which suggests that our findings are potentially relevant for current policy discussions. Further, regardless of the question, estimation of lifecycle effects requires the use of historical data. Second, we estimate ITT, not treatment-on-the-treated (TOT), effects. We note that an estimate of the TOT would also be informative, however, from a policy perspective, the ITT is perhaps useful as it is the policy, not other factors that affect insurance coverage, that is the lever that can be manipulated. Finally, the NLSY79 is a nationally representative data set, which can be troublesome for analyses of state-level treatment variables (Maclean, Tello-Trillo, and Webber 2019). However, we confirm our findings in the ASEC, which is representative at both the national and state level, and exclude states with poor coverage in the NLSY79, alleviating some of these concerns.

Our findings may be useful for thinking through how government regulation can distort labor market outcomes, both in the short-run as has been documented in other settings and, as we show in this study, over the lifecycle for affected groups of workers. While our study does not offer insight on the optimal set of insurance market regulations, it does imply that decision makers should consider the contemporaneous and persistent effects of such regulations, and heterogeneity across workers. Our study speaks to the labor literature that has documented the importance of aggregate demand shocks for lifecycle wages in that we show that additional shocks can lead to persistent effects, in particular, shocks from government regulation. We encourage more work, considering a broader set of shocks, to fully understand the challenges faced by young workers who are starting their careers and establishing lifecycle wage profiles.

Table 1. Labor market entrance cohort size among men: NLSY79 1979-2012

Labor market entrance year	Number of labor market entrants (unweighted)
1973	8
1974	57
1975	257
1976	468
1977	528
1978	600
1979	676
1980	534
1981	576
1982	510
1983	338
1984	201
1985	141
1986	98
1987	71
1988	34
1989	35
1990	20
1991	9
Total	5161

Notes: One observation per respondent in our analysis sample.

Table 2. State continuing coverage effective dates and months of mandated coverage

State	Effective date	Mandated months of coverage
Arkansas	7/20/1979	4
California	1/1/1985	3
Colorado	7/1/1986	3
Connecticut	10/1/1975	10
	1/1/1987	20
Georgia	7/1/1986	3
Illinois	1/1/1984	6
	8/23/1985	9
Kansas	1/1/1978	6
Kentucky	7/15/1980	9
Minnesota	8/1/1974	6
	3/19/1983	12
	6/1/1987	18
Missouri	9/28/1985	9
Nevada	1/1/1988	18
New Hampshire	8/22/1981	10
New Mexico	7/1/1983	6
New York	1/1/1986	3
North Carolina	1/1/1982	3
North Dakota	7/1/1983	10
Oklahoma	1/1/1976	1
Oregon	1/1/1982	6
Rhode Island	1988*	18
South Carolina	1/1/1979	2
South Dakota	7/1/1984	3
Tennessee	1/1/1981	3
Texas	1/1/1981	6
Utah	7/1/1986	2
Vermont	5/14/1986	6
Virginia	4/17/1986	3
Wisconsin	5/14/1980	18

Notes: Data source: Gruber and Madrian (1994).
*No effective day or month is provided. We impute January 1st, 1988 as the effective date for this state.

Table 3. Summary statistics among men: NLSY79 1979-2012

Variable:	**Mean/proportion**
Outcome variable	
Hourly wage (not log-transformed, 2012 dollars)	$22.38
Months of mandated continuing coverage	
Number of months (lagged on year)	1.020
Individual demographics	
Age at labor market entrance	18.93
White	0.822
African American	0.123
Hispanic	0.055
Less than high school at labor market entrance	0.144
High school at labor market entrance	0.514
Some college at labor market entrance	0.153
College degree at labor market entrance	0.188
Age-adjusted AFQT score	0.005
Mother’s education (years)	11.77
Father’s education (years)	12.01
Labor market entrance year	1980.0
Survey year	1992.9
Labor market entrance state level characteristics	
Female	0.514
Male	0.486
White	0.865
Non-white	0.136
0-18 years	0.321
19-64 years	0.567
65+ years	0.112
Per capita income (2012 dollars)	10,280
Employment-to-population ratio	0.749
Citizen ideology index (0-100)	47.08
Population	8,983,544
Count of high cost mandates	1.488
Observations	76694

Notes: The unit of observations is an NLSY79 respondent in a state in a year. NLSY79 weights applied.

Table 4. The contemporaneous effect of months of continuing coverage on the logarithm of constructed hourly wages among men: ASEC 1976-1992

Estimate:	**Beta coefficient (standard error)**
All men 18 to 54 years	
Sample mean hourly wage (not log-transformed)	$23.16
Months of continuing coverage	-0.0015* (0.0008)
Observations	517984
Men 18 to 30 years	
Sample mean hourly wage (not log-transformed)	$17.19
Months of continuing coverage	-0.0024** (0.0010)
Observations	221394
Men 31 to 54 years	
Sample mean hourly wage (not log-transformed)	$27.71
Months of continuing coverage	-0.0011 (0.0008)
Observations	296590

Notes: The unit of observations is an ASEC respondent in a state in a year. All models estimated with least squares and control for demographics, state fixed-effects, and year fixed-effects. ASEC sample weights applied. Standard errors are clustered around the state and reported in parentheses. ASEC wage data refers to the past calendar year, thus ASEC data 1976-1992 pertains to wages earned 1975-1991. Hourly wages are constructed from annual earnings from wages and salary, weeks worked, and usual hours worked per week.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 5A. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men: NLSY79 1979-2012

Sample:	**Beta coefficient (standard error)**
Sample mean hourly wage (not log-transformed)	$22.38
Months of continuing coverage × 1-5 years' experience	-0.0113***
	(0.0033)
Months of continuing coverage × 6-10 years' experience	0.0061***
	(0.0022)
Months of continuing coverage × 11-15 years' experience	0.0113***
	(0.0021)
Months of continuing coverage × 16-20 years' experience	0.0132***
	(0.0026)
Months of continuing coverage × 21-25 years' experience	0.0111***
	(0.0031)
Months of continuing coverage ×26-30 years' experience	0.0064**
	(0.0028)
Months of continuing coverage × 31+ years' experience	-0.0035
	(0.0055)
Observations	76694

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 5B. The persistent effect of months of continuing coverage at labor market entrance on the probability of a job change and the cumulative number of jobs held among men: NLSY79 1979-2012

Outcome:	**Job change in past year**	**Cumulative number of jobs**
Sample mean	0.369	8.322
Months of continuing coverage × 1-5 years' experience	0.0080*** (0.0018)	0.0444 (0.0349)
Months of continuing coverage × 6-10 years' experience	-0.0013 (0.0016)	0.0635** (0.0304)
Months of continuing coverage × 11-15 years' experience	0.0024 (0.0021)	0.0564 (0.0360)
Months of continuing coverage × 16-20 years' experience	-0.0013 (0.0017)	0.0479 (0.0366)
Months of continuing coverage × 21-25 years' experience	0.0014 (0.0019)	0.0618 (0.0483)
Months of continuing coverage ×26-30 years' experience	-0.0019 (0.0015)	0.0405 (0.0462)
Months of continuing coverage × 31+ years' experience	0.0027 (0.0057)	0.0533 (0.0773)
Observations	75440	76670

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with a linear probability model (job change) or least squares (cumulative number of jobs) and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 6A. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men by worker skill: NLSY79 1979-2012

Sample:	**Higher skill men**	**Lesser skill men**
Sample mean hourly wage (not log-transformed)	$29.47	$18.66
Months of continuing coverage × 1-5 years' experience	-0.0010	-0.0144***
	(0.0065)	(0.0030)
Months of continuing coverage × 6-10 years' experience	0.0109*	0.0011
	(0.0056)	(0.0030)
Months of continuing coverage × 11-15 years' experience	0.0147**	0.0048**
	(0.0056)	(0.0019)
Months of continuing coverage × 16-20 years' experience	0.0137***	0.0048
	(0.0049)	(0.0034)
Months of continuing coverage × 21-25 years' experience	0.0151**	-0.0002
	(0.0061)	(0.0029)
Months of continuing coverage ×26-30 years' experience	0.0085*	-0.0004
	(0.0050)	(0.0034)
Months of continuing coverage × 31+ years' experience	0.0055	-0.0061
	(0.0081)	(0.0068)
Observations	22077	53934

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses. Higher skill = enter the labor market with some college education. Lesser skill = enter the labor market with no college education.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 6B. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men by worker minority status: NLSY79 1979-2012

Sample:	**White men**	**Non-white men**
Sample mean hourly wage (not log-transformed)	$23.35	$17.87
Months of continuing coverage × 1-5 years' experience	-0.0092***	-0.0209***
	(0.0032)	(0.0069)
Months of continuing coverage × 6-10 years' experience	0.0076***	-0.0036
	(0.0021)	(0.0078)
Months of continuing coverage × 11-15 years' experience	0.0124***	0.0013
	(0.0023)	(0.0059)
Months of continuing coverage × 16-20 years' experience	0.0136***	0.0034
	(0.0027)	(0.0058)
Months of continuing coverage × 21-25 years' experience	0.0112***	0.0002
	(0.0032)	(0.0077)
Months of continuing coverage ×26-30 years' experience	0.0058*	0.0054
	(0.0030)	(0.0072)
Months of continuing coverage × 31+ years' experience	-0.0057	0.0089
	(0.0062)	(0.0067)
Observations	44330	32364

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 7. The persistent effect of months of continuing coverage at labor market entrance on ESI offer and labor supply outcomes among men: NLSY79 1979-2012

Outcome:	ESI offer[1]	Log (weeks)	Log (hours)	Log (earnings)
Sample proportion/mean (not log-transformed)	0.798	45.95	44.66	$46,083
Months of continuing coverage	-0.0075***	-0.0060***	-0.0059**	-0.0285***
× 1-5 years' experience	(0.0015)	(0.0015)	(0.0022)	(0.0039)
Months of continuing coverage	-0.0020	0.0019	0.0022*	0.0080
× 6-10 years' experience	(0.0015)	(0.0021)	(0.0012)	(0.0064)
Months of continuing coverage	-0.0016	0.0000	0.0023***	0.0124***
× 11-15 years' experience	(0.0019)	(0.0015)	(0.0008)	(0.0040)
Months of continuing coverage	-0.0017	-0.0003	0.0017	0.0170***
× 16-20 years' experience	(0.0018)	(0.0021)	(0.0011)	(0.0045)
Months of continuing coverage	-0.0022	0.0030**	0.0032**	0.0199***
× 21-25 years' experience	(0.0019)	(0.0014)	(0.0012)	(0.0041)
Months of continuing coverage	-0.0056***	-0.0011	0.0053***	0.0071
×26-30 years' experience	(0.0018)	(0.0016)	(0.0016)	(0.0053)
Months of continuing coverage	-0.0058	-0.0049	0.0042**	-0.0070
× 31+ years' experience	(0.0039)	(0.0045)	(0.0019)	(0.0064)
Observations	63791	74736	73851	71341

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with a linear probability model (binary outcome) or least squares (continuous outcome) and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.

[1]ESI=Employer-sponsored health insurance offer. This variable is not asked in the 1981 survey.

***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 8A. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of constructed hourly wages among men: ASEC 1976-2013

Estimate:	Beta coefficient (standard error)
Sample mean hourly wage (not log-transformed)	$23.89
Months of continuing coverage × 1-5 years' experience	-0.0092**
	(0.0036)
Months of continuing coverage × 6-10 years' experience	0.0037
	(0.0022)
Months of continuing coverage × 11-15 years' experience	0.0059***
	(0.0019)
Months of continuing coverage × 16-20 years' experience	0.0046**
	(0.0020)
Months of continuing coverage × 21-25 years' experience	0.0015
	(0.0017)
Months of continuing coverage × 26-30 years' experience	-0.0007
	(0.0020)
Months of continuing coverage × 31+ years' experience	-0.0066**
	(0.0026)
Observations	662546

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. ASEC sample weights applied. ASEC wage data pertain the previous calendar year, thus these data include wage data for 1975 to 2012. Hourly wages are constructed from annual earnings from wages and salary, weeks worked, and usual hours worked per week. Standard errors are clustered around the labor market entrance state and reported in parentheses. ***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 8B. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding states with less than 20 observations per labor market entrance year: NLSY79 1979-2012

Estimate:	Beta coefficient (standard error)
Sample mean hourly wage (not log-transformed)	$22.34
Months of continuing coverage × 1-5 years' experience	-0.0123***
	(0.0036)
Months of continuing coverage × 6-10 years' experience	0.0059**
	(0.0023)
Months of continuing coverage × 11-15 years' experience	0.0109***
	(0.0021)
Months of continuing coverage × 16-20 years' experience	0.0130***
	(0.0027)
Months of continuing coverage × 21-25 years' experience	0.0100***
	(0.0032)
Months of continuing coverage × 26-30 years' experience	0.0060**
	(0.0027)
Months of continuing coverage × 31+ years' experience	-0.0037
	(0.0056)
Observations	76338

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses. ***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 9. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men accounting for Federal COBRA: NLSY79 1979-2012

Model:	**Model (1)**	**Model (2)**
Sample mean hourly wage (not log-transformed)	22.13	22.38
Months of continuing coverage	-0.0157***	-0.0147***
× 1-5 years' experience	(0.0019)	(0.0028)
Months of continuing coverage	0.0021	-0.0000
× 6-10 years' experience	(0.0035)	(0.0030)
Months of continuing coverage	0.0082**	0.0077***
× 11-15 years' experience	(0.0033)	(0.0027)
Months of continuing coverage	0.0127***	0.0109***
× 16-20 years' experience	(0.0038)	(0.0026)
Months of continuing coverage	0.0102**	0.0118***
× 21-25 years' experience	(0.0041)	(0.0040)
Months of continuing coverage	0.0087*	0.0082*
× 26-30 years' experience	(0.0044)	(0.0043)
Months of continuing coverage	-0.0073	-0.0080
× 31+ years' experience	(0.0066)	(0.0065)
Observations	75634	76694

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
Model (1): Exclude respondents who entered the labor market after 1987.
Model (2): Assign 9 months of mandated continuing coverage in 1986 and 18 months of mandated continuing coverage in 1987 to 1991 to all states with the exception of Connecticut which mandated 20 months 1986 to 1991.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 10A. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of wages among men using alternative specifications: NLSY79 1979-2012

Sample:	**Model (1)**	**Model (2)**	**Model (3)**	**Model (4)**	**Model (5)**
Sample mean hourly wage (not log-transformed)	$22.38	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage ×					
1-5 years' experience	-0.0129***	-0.0114***	-0.0142***	-0.0127***	-0.0126***
	(0.0038)	(0.0039)	(0.0027)	(0.0039)	(0.0037)
6-10 years' experience	0.0029	0.0060**	0.0032	0.0012	0.0044*
	(0.0025)	(0.0025)	(0.0037)	(0.0021)	(0.0023)
11-15 years' experience	0.0078***	0.0110***	0.0085**	0.0075***	0.0099***
	(0.0024)	(0.0023)	(0.0042)	(0.0020)	(0.0019)
16-20 years' experience	0.0102***	0.0128***	0.0102**	0.0088***	0.0112***
	(0.0028)	(0.0026)	(0.0047)	(0.0026)	(0.0021)
21-25 years' experience	0.0071**	0.0105***	0.0081	0.0053*	0.0104***
	(0.0031)	(0.0033)	(0.0050)	(0.0028)	(0.0029)
26-30 years' experience	0.0025	0.0062**	0.0040	0.0032	0.0056**
	(0.0029)	(0.0028)	(0.0051)	(0.0029)	(0.0026)
31+ years' experience	-0.0091	-0.0038	-0.0060	-0.0048	-0.0049
	(0.0061)	(0.0048)	(0.0059)	(0.0047)	(0.0051)
Observations	76694	76694	76694	76694	76694

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects unless otherwise noted. NLSY79 sample weights applied unless otherwise noted. Standard errors are clustered around the labor market entrance state and reported in parentheses.
Model (1): Exclude labor market entrance state-level time-varying and individual-level variables.
Model (2): Include labor market entrance division-by-year fixed-effects.
Model (3): Include labor market entrance state-specific linear time trends.
Model (4): Do not lag mandated months of continuing coverage by one year.
Model (5): Include survey state fixed-effects.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 10B. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of wages among men using alternative weighting schemes and samples: NLSY79 1979-2012

Sample:	**Model (1)**	**Model (2)**
Sample mean hourly wage (not log-transformed)	$20.34┼	$22.52
Months of continuing coverage ×		
1-5 years' experience	-0.0147***	-0.0119***
	(0.0039)	(0.0031)
6-10 years' experience	0.0030	0.0048*
	(0.0030)	(0.0027)
11-15 years' experience	0.0081***	0.0098***
	(0.0019)	(0.0022)
16-20 years' experience	0.0111***	0.0114***
	(0.0024)	(0.0024)
21-25 years' experience	0.0102***	0.0092***
	(0.0030)	(0.0030)
26-30 years' experience	0.0057**	0.0045
	(0.0023)	(0.0027)
31+ years' experience	-0.0014	-0.0054
	(0.0047)	(0.0048)
Observations	76694	72371

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects unless otherwise noted. NLSY79 sample weights applied unless otherwise noted. Standard errors are clustered around the labor market entrance state and reported in parentheses.
Model (1): Remove survey weights.
Model (2): Exclude respondents who entered the labor market before 1976.
┼Mean value is unweighted.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 10C. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of wages among men using alternative imputation approaches for labor market state: NLSY79 1979-2012

Sample:	**Model (1)**	**Model (2)**
Sample mean hourly wage (not log-transformed)┼	$22.39	$22.39
Months of continuing coverage ×		
1-5 years' experience	-0.0118***	-0.0099***
	(0.0042)	(0.0032)
6-10 years' experience	0.0058**	0.0077***
	(0.0024)	(0.0022)
11-15 years' experience	0.0117***	0.0139***
	(0.0017)	(0.0024)
16-20 years' experience	0.0130***	0.0152***
	(0.0018)	(0.0030)
21-25 years' experience	0.0149***	0.0168***
	(0.0029)	(0.0053)
26-30 years' experience	0.0084***	0.0102**
	(0.0025)	(0.0044)
31+ years' experience	-0.0110	-0.0087
	(0.0070)	(0.0053)
Observations	76694	76099

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects unless otherwise noted. NLSY79 sample weights applied unless otherwise noted. Standard errors are clustered around the labor market entrance state and reported in parentheses.
Model (1): Assign state of residence in 1979 as the labor market entrances state.
Model (2): Assign the birth state as the labor market entrances state.
┼Mean value is unweighted. Sample sizes differ across Model (1) and Model (2) due to differences in missingness in state of residence in 1979 and birth state. Details available on request from the corresponding author.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11A. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**AR**	**CA**	**CO**	**CT**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0110***	-0.0114***	-0.0113***	-0.0114**
× 1-5 years' experience	(0.0033)	(0.0034)	(0.0033)	(0.0043)
Months of continuing coverage	0.0061***	0.0058**	0.0058***	0.0053**
× 6-10 years' experience	(0.0022)	(0.0023)	(0.0021)	(0.0026)
Months of continuing coverage	0.0111***	0.0109***	0.0109***	0.0123***
× 11-15 years' experience	(0.0020)	(0.0020)	(0.0020)	(0.0027)
Months of continuing coverage	0.0132***	0.0128***	0.0130***	0.0151***
× 16-20 years' experience	(0.0026)	(0.0025)	(0.0026)	(0.0034)
Months of continuing coverage	0.0110***	0.0103***	0.0109***	0.0130***
× 21-25 years' experience	(0.0031)	(0.0028)	(0.0031)	(0.0041)
Months of continuing coverage	0.0066**	0.0065**	0.0060**	0.0077**
× 26-30 years' experience	(0.0028)	(0.0027)	(0.0028)	(0.0037)
Months of continuing coverage	-0.0036	-0.0039	-0.0042	0.0028
× 31+ years' experience	(0.0055)	(0.0055)	(0.0056)	(0.0044)
Observations	75866	69192	75224	75081

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11B. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**GA**	**IL**	**IA**	**KS**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0111***	-0.0105***	-0.0111***	-0.0108***
× 1-5 years' experience	(0.0033)	(0.0030)	(0.0032)	(0.0033)
Months of continuing coverage	0.0062***	0.0066***	0.0057**	0.0063***
× 6-10 years' experience	(0.0022)	(0.0021)	(0.0021)	(0.0022)
Months of continuing coverage	0.0115***	0.0119***	0.0110***	0.0111***
× 11-15 years' experience	(0.0022)	(0.0022)	(0.0021)	(0.0021)
Months of continuing coverage	0.0133***	0.0136***	0.0133***	0.0124***
× 16-20 years' experience	(0.0026)	(0.0026)	(0.0028)	(0.0022)
Months of continuing coverage	0.0110***	0.0118***	0.0103***	0.0105***
× 21-25 years' experience	(0.0031)	(0.0032)	(0.0029)	(0.0029)
Months of continuing coverage	0.0062**	0.0065**	0.0064**	0.0057**
× 26-30 years' experience	(0.0028)	(0.0028)	(0.0029)	(0.0025)
Months of continuing coverage	-0.0038	-0.0034	-0.0033	-0.0058
× 31+ years' experience	(0.0055)	(0.0055)	(0.0055)	(0.0050)
Observations	73438	74028	75809	76227

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11C. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**KY**	**MN**	**MO**	**NV**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0113***	-0.0109***	-0.0113***	-0.0112***
× 1-5 years' experience	(0.0033)	(0.0037)	(0.0034)	(0.0033)
Months of continuing coverage	0.0060***	0.0052**	0.0058***	0.0062***
× 6-10 years' experience	(0.0022)	(0.0022)	(0.0021)	(0.0022)
Months of continuing coverage	0.0113***	0.0106***	0.0111***	0.0115***
× 11-15 years' experience	(0.0021)	(0.0018)	(0.0020)	(0.0021)
Months of continuing coverage	0.0132***	0.0135***	0.0128***	0.0133***
× 16-20 years' experience	(0.0026)	(0.0030)	(0.0025)	(0.0026)
Months of continuing coverage	0.0111***	0.0109***	0.0108***	0.0111***
× 21-25 years' experience	(0.0031)	(0.0034)	(0.0030)	(0.0031)
Months of continuing coverage	0.0064**	0.0063**	0.0060**	0.0064**
× 26-30 years' experience	(0.0028)	(0.0030)	(0.0027)	(0.0028)
Months of continuing coverage	-0.0036	-0.0035	-0.0038	-0.0036
× 31+ years' experience	(0.0055)	(0.0067)	(0.0056)	(0.0055)
Observations	76594	74701	74795	76573

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11D. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**NH**	**NM**	**NY**	**NC**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0113***	-0.0112***	-0.0117***	-0.0112***
× 1-5 years' experience	(0.0033)	(0.0034)	(0.0034)	(0.0034)
Months of continuing coverage	0.0061***	0.0061***	0.0060***	0.0062***
× 6-10 years' experience	(0.0022)	(0.0021)	(0.0022)	(0.0022)
Months of continuing coverage	0.0113***	0.0113***	0.0110***	0.0114***
× 11-15 years' experience	(0.0021)	(0.0021)	(0.0020)	(0.0021)
Months of continuing coverage	0.0132***	0.0132***	0.0130***	0.0132***
× 16-20 years' experience	(0.0026)	(0.0026)	(0.0026)	(0.0026)
Months of continuing coverage	0.0111***	0.0110***	0.0105***	0.0109***
× 21-25 years' experience	(0.0031)	(0.0030)	(0.0030)	(0.0031)
Months of continuing coverage	0.0064**	0.0063**	0.0059**	0.0062**
× 26-30 years' experience	(0.0028)	(0.0028)	(0.0027)	(0.0028)
Months of continuing coverage	-0.0035	-0.0035	-0.0037	-0.0038
× 31+ years' experience	(0.0055)	(0.0055)	(0.0056)	(0.0056)
Observations	76585	75872	71140	73716

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11E. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**ND**	**OK**	**OR**	**RI**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0113***	-0.0110***	-0.0109***	-0.0113***
× 1-5 years' experience	(0.0034)	(0.0032)	(0.0033)	(0.0033)
Months of continuing coverage	0.0060***	0.0062***	0.0066***	0.0061***
× 6-10 years' experience	(0.0022)	(0.0021)	(0.0021)	(0.0022)
Months of continuing coverage	0.0113***	0.0116***	0.0117***	0.0113***
× 11-15 years' experience	(0.0021)	(0.0022)	(0.0021)	(0.0021)
Months of continuing coverage	0.0131***	0.0135***	0.0128***	0.0132***
× 16-20 years' experience	(0.0026)	(0.0027)	(0.0024)	(0.0026)
Months of continuing coverage	0.0110***	0.0113***	0.0108***	0.0111***
× 21-25 years' experience	(0.0031)	(0.0032)	(0.0029)	(0.0031)
Months of continuing coverage	0.0063**	0.0066**	0.0056**	0.0064**
× 26-30 years' experience	(0.0028)	(0.0029)	(0.0023)	(0.0028)
Months of continuing coverage	-0.0036	-0.0031	-0.0034	-0.0035
× 31+ years' experience	(0.0055)	(0.0055)	(0.0056)	(0.0055)
Observations	76577	75652	76340	76691

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11F. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**SC**	**SD**	**TN**	**TX**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage	-0.0111***	-0.0114***	-0.0114***	-0.0101***
× 1-5 years' experience	(0.0033)	(0.0034)	(0.0033)	(0.0032)
Months of continuing coverage	0.0061***	0.0061***	0.0059***	0.0076***
× 6-10 years' experience	(0.0023)	(0.0022)	(0.0022)	(0.0018)
Months of continuing coverage	0.0112***	0.0113***	0.0112***	0.0116***
× 11-15 years' experience	(0.0021)	(0.0021)	(0.0021)	(0.0022)
Months of continuing coverage	0.0131***	0.0132***	0.0131***	0.0130***
× 16-20 years' experience	(0.0026)	(0.0026)	(0.0026)	(0.0025)
Months of continuing coverage	0.0108***	0.0111***	0.0109***	0.0122***
× 21-25 years' experience	(0.0030)	(0.0031)	(0.0031)	(0.0034)
Months of continuing coverage	0.0066**	0.0064**	0.0063**	0.0069**
× 26-30 years' experience	(0.0029)	(0.0028)	(0.0028)	(0.0028)
Months of continuing coverage	-0.0037	-0.0035	-0.0035	-0.0029
× 31+ years' experience	(0.0054)	(0.0055)	(0.0056)	(0.0055)
Observations	74741	76489	75596	71183

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 11G. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men excluding treatment states one at a time: NLSY79 1979-2012

Exclude:	**UT**	**VT**	**VA**	**WI**
Sample mean hourly wage (not log-transformed)┼	$22.38	$22.38	$22.38	$22.38
Months of continuing coverage × 1-5 years' experience	-0.0113***	-0.0113***	-0.0111***	-0.0172***
	(0.0033)	(0.0033)	(0.0032)	(0.0057)
Months of continuing coverage × 6-10 years' experience	0.0060***	0.0061***	0.0062***	0.0056
	(0.0022)	(0.0022)	(0.0022)	(0.0057)
Months of continuing coverage × 11-15 years' experience	0.0113***	0.0113***	0.0113***	0.0124**
	(0.0021)	(0.0021)	(0.0022)	(0.0060)
Months of continuing coverage × 16-20 years' experience	0.0132***	0.0132***	0.0131***	0.0148**
	(0.0026)	(0.0026)	(0.0026)	(0.0073)
Months of continuing coverage × 21-25 years' experience	0.0111***	0.0111***	0.0109***	0.0133*
	(0.0031)	(0.0031)	(0.0031)	(0.0078)
Months of continuing coverage × 26-30 years' experience	0.0064**	0.0064**	0.0062**	0.0093
	(0.0028)	(0.0028)	(0.0028)	(0.0078)
Months of continuing coverage × 31+ years' experience	-0.0036	-0.0035	-0.0038	-0.0063
	(0.0055)	(0.0055)	(0.0055)	(0.0104)
Observations	76590	76544	74777	73761

Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Sample means are based on the full sample of states.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 12. Correlates of months of mandated continuing coverage: 1973-1990

Outcome:	**Months of mandated continuing coverage**
Sample mean months of mandated continuing coverage	1.44
Female	-204.7058
	(202.5292)
Non-white	-13.6376
	(15.9725)
18 to 64 years	56.3720
	(90.3117)
65+ years	-140.0522**
	(68.2904)
Income per capita (2012 dollars)	-0.0000
	(0.0003)
Employment-to-population ratio	2.5393
	(10.3022)
Citizen ideology index (0-100)	-0.0451
	(0.0496)
Population	-0.0000
	(0.0000)
Count of high cost mandates	1.3218
	(0.8465)
F-statistic of covariate joint significance (*p*-value)	1.87
	(0.0879)
Observations	1008

Notes: The unit of observations is a state in a year. All models estimated with least squares and control for state fixed-effects, and year fixed-effects. Omitted categories are male, white, and ages 0 to 17 years. Data are weighted by the state population. Standard errors are clustered around the state and reported in parentheses.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 13A. The effect of months of mandated continuing coverage of the composition of labor market entrants: NLSY79 1979-2012

Outcome:	**White**	**Rural**	**Born outside the U.S.**	**Mother's education**	**Father's education**
Sample mean/proportion	0.815	0.225	0.033	11.78	12.02
Months of mandated continuing coverage	0.0005 (0.0019)	0.0026 (0.0021)	-0.0020 (0.0014)	-0.0051 (0.0176)	0.0244 (0.0170)
Observations	5161	5161	5161	5161	5161

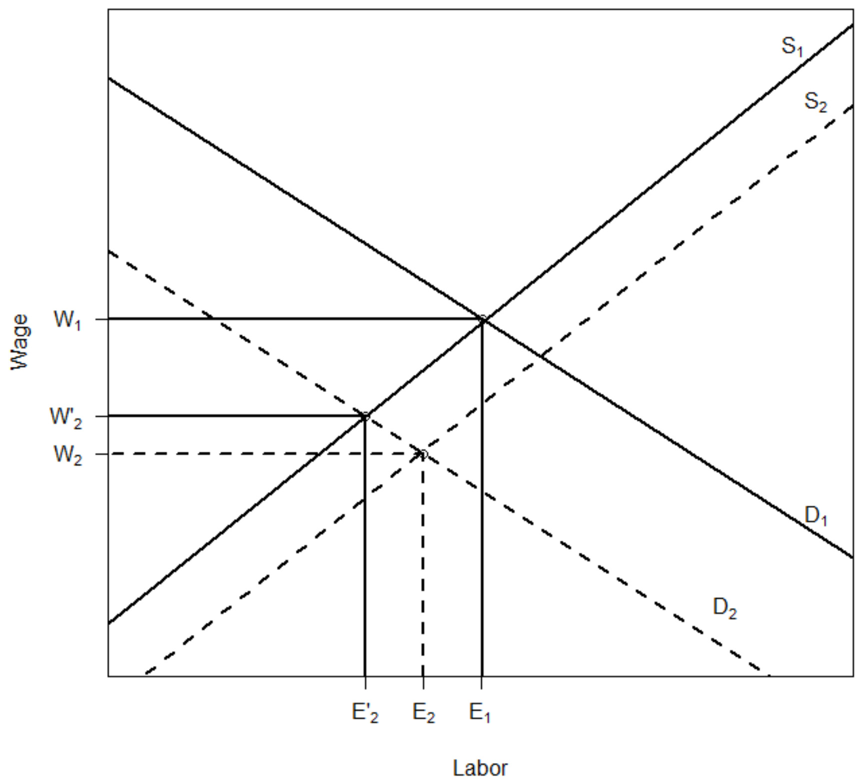
Notes: The unit of observations is an NLSY79 respondent in a state in a year. The sample size is smaller than the main analysis sample as we use just one observation per respondent. All models estimated with a linear probability model (binary outcomes) or least squares (continuous outcomes) and control for individual demographics, labor market entrance state characteristics, labor market entrance state fixed-effects, and labor market entrance year fixed-effects. NLSY79 sample weights applied. Standard errors are clustered around the labor market entrance state and reported in parentheses.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Table 13B. The persistent effect of months of continuing coverage at labor market entrance on the logarithm of wages among men controlling for additional individual-level variables: NLSY79 1979-2012

Sample mean hourly wage (not log-transformed)	**$22.13**
Months of continuing coverage ×	
1-5 years' experience	-0.0149*** (0.0029)
6-10 years' experience	-0.0000 (0.0027)
11-15 years' experience	0.0070*** (0.0023)
16-20 years' experience	0.0096*** (0.0024)
21-25 years' experience	0.0101** (0.0038)
26-30 years' experience	0.0059 (0.0038)
31+ years' experience	-0.0088 (0.0046)
Observations	76694

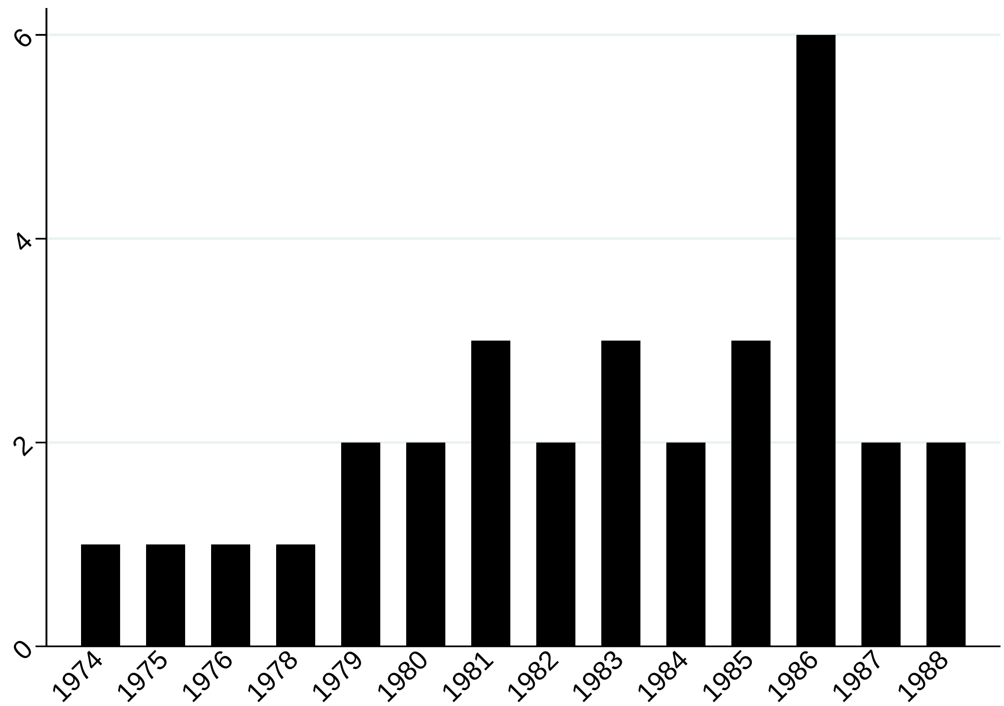
Notes: The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, additional individual level controls (see outcomes in Table 13A), labor market entrance state characteristics, labor market entrance state fixed-effects, labor market entrance year fixed-effects, and survey year fixed-effects unless otherwise noted. NLSY79 sample weights applied unless otherwise noted. Standard errors are clustered around the labor market entrance state and reported in parentheses.
┼Mean value is unweighted.
***;**;*=statistically different from zero at the 1%; 5%;10% level.

Figure 1. The effect of mandated benefits on wages and employment



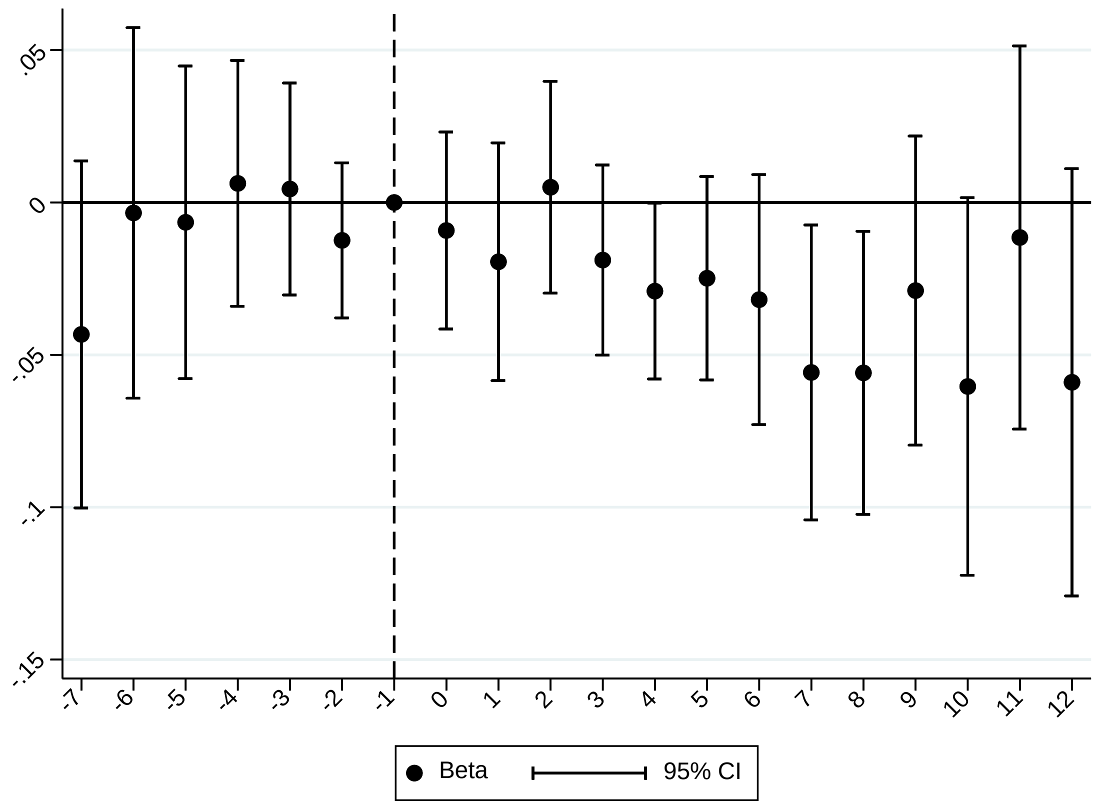
Notes: Figure based on Summers (1989). The magnitude of the shifts in the demand and supply curves are arbitrarily chosen and are for illustrative purposes only.

Figure 2. Number of state-level continuing coverage mandate changes by labor market entrance year:



Notes: Source: Gruber and Madrian (1994). See Table 2 for effective dates.

Figure 3. The effect of months of continuing coverage at labor market entrance on the logarithm of hourly wages among men using an event study: NLSY79 1979-2012



Notes: Sample mean average wage (not log transformed) is $ 16.71. *N*= 53466. The unit of observations is an NLSY79 respondent in a state in a year. All models estimated with least squares and control for individual demographics, state characteristics, state fixed-effects, and year fixed-effects. NLSY79 sample weights applied. 95% confidence intervals account for within-state clustering and are reported with vertical lines. The omitted category is the year prior to the mandate's effective date (-1). See text for details.

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