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Who Cares? Paid Sick Leave Mandates, Care-Giving, and Gender

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ABSTRACT

We use employment data from the Current Population Survey to assess the efficacy of state-mandated paid sick leave policies on leave-taking behavior with a focus on any variation by gender. We find that these policies increase leave taking for care-giving for men by 10-20%, and this effect is strongest for men with young children in the household. In addition, we find that Hispanic men and men without a bachelor's degree, who historically have had low access to paid sick leave, are 20–25% more likely to take care-giving leave. By comparison, we do not find evidence that these policies affect leave taking for own sickness for men or women, nor do we find evidence that these policies affect care-giving leave taking for women. Our evidence highlights the importance of studying care-giving leave within the context of paid leave policies and the importance of considering gender differences in the treatment effect within this context.

Keywords: Paid Sick Leave, Care-Giving Leave, Gender, Leave Taking
JEL Classification: H31; J14; J38

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1. Introduction

In 1950, gender-specific specialization in household and market production was standard. On average, women spent almost 5 times as many hours in home production as men, and men worked for pay almost 3.5 times as many hours as women (Ramey, 2009). Since the 1950s, however, female labor force participation in the U.S. increased dramatically before leveling off in the 1990s. Participation among women aged 25–49 increased from roughly 30% to 75% between 1950 and 1990, and increased participation of mothers of young children was a major factor behind this historic shift (Ruggles et al., 2023). The gender gap in earnings, which was partially driven by women's career interruptions around childbearing, also predictably declined through the 1990s before leveling off (Blau and Kahn, 2017). At the same time, gender gaps in home production narrowed as both men and women increasingly split their time across home and market production (see Figure 1).

Despite the large gains made by women in labor market, substantial gender disparities remain in both unpaid household work and market earnings. Women spend twice as many hours as men doing housework (Ramey, 2009) and employed fathers spend roughly half as much time caring for their children as employed mothers (Guryan, Hurst and Kearney, 2008). Moreover, employed women are three times as likely to take on the responsibility for household children's health and four times more likely than fathers to take time off from work to care for a sick child (Gomez, Ranji and Salganicoff, 2018). The stagnation of the earnings gap since the 1990's has become an oft-debated puzzle in recent literature and studies of the "child penalty" have risen in prominence (Kleven et al., 2019; Lurie, Patel and Ramnath, 2022; Byker, 2016; Yu and Hara, 2021). The broad conclusion from this strand of literature is that the onset of parenthood leads to a disruption in labor market outcomes for mothers, but not for fathers.

Family friendly policies are frequently proposed as a solution for the child penalty women face in the labor market. However, a growing literature is finding that gender-neutral polices, ranging from university tenure clock pauses to state and national paid parental leave policies, have not reduced the gender gap and may even be detrimental to women's relative career trajectories (Antecol, Bedard and Stearns, 2018; Bailey et al., 2019; Kleven et al., 2020). Still, the role of paid sick leave policies that mandate pay for short absences from work due to own or a family member's illness is less understood in this context.

In this paper, we estimate gender differences in the treatment effect of state-mandated access to paid sick leave on leave-taking used for own sickness and for care-giving. Historically, American workers have relied on the voluntary provision of paid sick leave benefits by employers, and, as a result, there are widespread systematic differences in access. In response, some state and local governments passed laws that require employees within their jurisdiction to accrue paid sick leave for use on one's own illness or to care for a sick family member. Connecticut was the first state to implement such a policy in 2012, and, to date, sixteen additional states have passed similar laws.

Our analysis compares leave-taking behavior by individuals living in treated states (states that pass mandates) to individuals living in untreated states (states that do not pass mandates during our analysis window). Our empirical identification strategy is based on a two-way fixed effects model, which allows for the staggered treatment assignment driven by the variation in the timing of when states passed their mandates. Within this context, we account for the well-documented biases of these models by implementing a two-stage differences-in-differences estimator of the average treatment effect (Goodman-Bacon, 2021; Gardner, 2022). Ultimately, we estimate the effect of paid sick leave mandates that were passed in D.C. and in ten states through the end of 2018 using monthly employment data captured by the Current Population Survey.[1]

Our results paint an updated and nuanced picture of the treatment effects of mandated access to paid sick leave. In aggregate, we find that paid sick leave has little impact on the likelihood of work absences for either own illness or for care-giving. However, these results mask stark heterogeneity by gender for leave used to accommodate care-giving. When we separately estimate results for men and women, we find that paid sick leave policies increase the likelihood that men take care-giving leave by 10-20%, and these results are driven by men with children in the household. In addition, we find strong effects of the policy on those men who historically had the least access to paid sick leave through their employer: Hispanic men and men with less than a bachelor's degree. By comparison, we find no effect of the same mandates on the likelihood of a work absence due to care-giving for women, regardless of whether or not there are children in the household, an individual's race and ethnicity, or an individual's educational attainment.

Proposals for federally mandated access to paid sick leave have been an active source of political discourse during the last two decades. In 2005, Senator Theodore Kennedy introduced the Health Families Act in the 108th Congress, which was the the first legislative initiative for a federal paid sick leave mandate. The Healthy Families Act has been reintroduced in every Congress since 2005, but the bill has failed to gain the support of Congress. In 2015, President Barack Obama became the first president to call for paid leave during his State of the Union address:

> Today, we are the only advanced country on Earth that doesn't guarantee paid sick leave or paid maternity leave to our workers. Forty-three million workers have no

[1] We do not study the effects of mandates that were implemented in 2019 and beyond to avoid the confounding influence of the pandemic on leave-taking during the post-treatment period. See Table A.1 for a detailed description of the paid sick leave mandates that passed in each state.

> paid sick leave – 43 million. Think about that. And that forces too many parents to make the gut-wrenching choice between a paycheck and a sick kid at home. So I'll be taking new action to help states adopt paid leave laws of their own. And since paid sick leave won where it was on the ballot last November, let's put it to a vote right here in Washington. Send me a bill that gives every worker in America the opportunity to earn seven days of paid sick leave. It's the right thing to do.
> *President Barack Obama, State of the Union, January 20, 2015*

In that same year, President Obama signed an executive order requiring that employees of federal contractors accrue a minimum of one hour of paid sick leave for every 30 hours worked, with a maximum accrual amount of no less than 56 hours per year.

Our results contribute to these ongoing policy debates on broader provision of sick leave. We find that paid sick leave mandates have reach beyond what had previously been estimated. Early studies of the first paid sick leave mandates found mixed evidence on whether mandating access affects leave-taking behavior (Stearns and White, 2018; Maclean, Pichler and Ziebarth, 2020) or employment and earnings (Callison and Pesko (2016); Ahn and Yelowitz (2016); Pichler and Ziebarth (2020); Al-Sabah and Ouimet (2022)). Other studies found evidence that mandating paid sick leave improved health (Stearns and White, 2018; Pichler and Ziebarth, 2017) and increased health care utilization (DeRigne, Stoddard-Dare and Quinn, 2016; Jeung, Lee and Gimm, 2021). However, these studies focused on the effect of mandated paid sick leave on work absences related to own illness. Moreover, the generalizability of these studies is limited by the fact that they are based on evidence drawn from early adopters, D.C. and Connecticut, which both constitute small populations where the workforce is not necessarily representative of the broader U.S. labor market.

Our work compliments and expands this existing literature on paid sick leave in several important ways. First, our analysis brings new and more generalizable evidence to the policy debate by studying the effect of mandates passed in nine additional states that have not yet been investigated. Second, our analysis is the first to specifically focus on the effect of mandated paid sick leave on care-giving work absences. Our evidence provides a possible explanation for why past literature finds limited effects of these mandates on leave taking: previous studies focus on leaves for sickness rather than for care-giving. Finally, our evidence highlights important heterogeneity by gender; we find that mandated access to paid sick leave nudges men, and especially men with young children, into additional care-giving. To the extent that women are more likely to take care-giving relative to men, these policies can help close the gender gap in leave-taking to care for a sick family member. As such, these findings contribute to a broader conversation about gender norms and women's labor market outcomes.

2. Background

2.1. Paid Sick Leave in the U.S.

The Family and Medical Leave Act (FMLA), passed in 1993, is the only federal regulation governing the provision of leave benefits for U.S. workers. However, its provisions, while historic, are minimal. FMLA guarantees job protection for eligible employees of certain employers who take up to 12 weeks of *unpaid leave* for work absences due to own illness or injury, to care for an ill or injured family member, or to care for a newborn or newly adopted child.[2] There are no federal regulations requiring that private employers provide *paid* leave for either short- or long-duration work absences.

There is an important distinction to be made between paid sick leave policies, which relate to short absences and are the topic of this paper, and state and federal efforts to expand access to paid Family and Medical leave, which have also been a source of vigorous policy debate and are outside the scope of this analysis. California was the first state to implement a paid family leave policy in 2004, and, to date, 11 states (including D.C.) have implemented similar policies.[3] Reasons for work absences under paid family and medical leave policies can overlap with reasons for work absences taken under paid sick leave — namely absences due to own and family illness or injury. However, the expected duration of leave taken under the two policies differs substantially: paid family and medical leave policies are typically designed to provide employment insurance for medium-to-long duration work absences — leaves of 1 to 12 weeks — whereas paid sick leave policies are designed for short-duration work absences — leaves of a few hours and up to 7 to 10 days. Consistent with this, many state paid family and medical leave policies include a waiting period of one week or more; only work absences that extend beyond this period are eligible for wage replacement. In addition, paid family and medical leave is typically tied to a payroll tax that funds a trust used to pay benefits, or wage replacement, to eligible employees. By comparison, paid sick leave policies represent an unfunded liability to employers, who must pay benefits directly to employees.

Absent any national policy for paid *sick* leave, some state and local governments began regulating access to paid sick leave for employees who work in their jurisdiction.[4] In 2008,

[2] Employees are eligible for leave under FMLA if they have worked for their employer at least 12 months, at least 1,250 hours over the past 12 months, and work at a location where the company employs 50 or more employees within 75 miles.

[3] The following states have implemented paid family and medical leave laws: California, Colorado, Connecticut, D.C., Maryland, Massachusetts, New Jersey, New York, Oregon, Rhode Island, and Washington.

[4] Prior to 2008, a handful of municipalities implemented locality-specific paid sick leave mandates.

Washington D.C. became the first large government to mandate that qualified employees accrue one hour of paid sick leave for every 43 hours worked, up to a maximum of 7 days per year. Connecticut followed suit in 2012 as the first state to mandate paid sick leave; in this case, full-time service sector employees at large firms must accrue one hour of sick leave for every 40 hours worked, up to a maximum of 40 hours per year. Between 2014 and 2018, ten more states adopted paid sick leave mandates with even more generous benefits, and in all cases capped the total annual paid sick leave earned to roughly one week of leave. Table A.1 summarizes the timing and policy details of these mandates, including an estimate of the population of workers likely affected. All of these enacted mandates allow employees to take leave due to either an employee's or a family member's illness, injury or condition, or preventative care. In other words, these mandates allow employees to take *short-duration* leave for what is typically termed as either sick leave or care-giving leave.

Regardless of federal policy, a large fraction of employers voluntarily provide paid sick leave benefits to at least a subset of their employees. According to the National Health Interview Survey (NHIS), in 2010—two years prior to the first state mandate—roughly 59% of the private-sector workforce aged 25 to 49 had paid sick leave benefits through their employer.[5] However, these figures mask the considerable heterogeneity in access to paid sick leave across individuals. In that same year, there was a 25 percentage point gap in access between workers at small (those with fewer than 50 employees) and large employers (those with 50 or more employees): 46% of workers at small employers had access to paid leave compared with 71%

[5]There are two primary public-use data sources that report on access to paid sick leave: the Bureau of Labor Statistics National Compensation Survey (NCS) and the NHIS. For comparison purposes, the NCS reports that roughly 63% of the private-sector workforce (with no age restriction) had access to paid sick leave benefits through their employer in 2010. The NCS is conducted at the establishment-by-job level and contains little demographic information about employees; importantly, the NCS does not report statistics for male and female employees separately. By comparison, the NHIS reports whether an individual has paid sick leave through their employer in addition to detailed individual characteristics, including gender and parenthood status as well as characteristics of the job including firm size and hours worked. For this reason, and to ensure consistency across all reported statistics in this background section, we calculate statistics based on the NHIS for the civilian population employed in the private sector aged 25–49 using data from 2006–2018.

at large employers. Part time employees were also 39 percentage points less likely to have paid sick leave benefits (27%, compared to 66% of full-time employees).[6] Finally, less than 20% of construction workers and food service workers had access to paid sick leave compared to the 86% of professionals in STEM industries and 78% of those employed in social work.

2.2. Unequal Access to Paid Sick Leave in the U.S.

The disparities in paid sick leave access over time are more clearly visualized in Figure 2. Figure 2 reports trends in the likelihood of access to paid sick leave from 2006–2018 separately for men and women by pooling information across NHIS survey rounds for private-sector workers aged 25–49. Panel (a) reveals two facts. First, access to paid sick leave was stable from 2006 through 2014, where roughly 58% of workers had access through their employers. After 2012, when the first state implemented a paid sick leave mandate, access slowly increased, landing at 66% by 2017. Second, while women are more likely than men to have access to paid sick leave prior to 2012, the gender gap is qualitatively small and narrowed after 2012. In 2012, for example, 57% of men and 60% of women had access to paid sick leave; by 2017, 66% of men and 67% of women had access to paid sick leave. Panels (b)-(d), however, show that very similar aggregate levels of access by gender in truth conceal important differences by gender related to individual earnings, full-time/part-time status, and family characteristics.

Panel (b) illustrates access to paid sick leave for men and women based on whether they are low income, where low income individuals are defined as those earning less than $35,000 per year. Low income workers are less likely to have access to paid sick leave, and women and men are differentially likely to be low income workers. By comparing panels (a) and (b),

[6] The NHIS does not report the number of days of paid sick leave. The 2010 NCS estimates that workers with access to paid sick leave and one to five years of tenure with a firm had, on average, nine paid sick days.

we see that equal access to paid leave between men and women on average masks substantial heterogeneity in access by gender once we decompose this statistic by earnings. Prior to the first state policy in 2012, women were more likely to have access to paid sick leave than men: the gender-gap in access was around 10 percentage points for those earning less than $35,000 and for those earning more than $35,000. This gender gap likely reflects a stronger preference among women for family-friendly jobs that offer these types of benefits (Goldin, 2014). As states began implementing mandated access to paid sick leave, the gender-gap in access shrinks by 50% among low-income workers, whereas the gap is virtually unchanged among individuals earning more than $35,000. Said differently, in the years coinciding with expanded access to mandated sick leave, low-wage men make gains in their relative likelihood of access to paid sick leave compared to low-wage women.

Panel (c), which depicts access to paid sick leave by full-time and part time employment, highlights a notable difference in access by gender in full-time positions (and a less pronounced difference for part time workers). Among full-time workers, women are employed in jobs with considerably higher rates of paid sick leave access than men – again, potentially driven by women seeking out more family-friendly workplaces. An alternative way to achieve schedule flexibility to accommodate family care responsibilities is to work part time, but part-time jobs are much less likely to provide paid sick leave. These compositional differences in job types by gender result in a relatively narrow gender gap overall despite the wide female advantage in access amongst full-time workers. For example, in 2010 27% of women were employed in part time work compared with just 11% of men.

Finally, panel (d) reveals pronounced gender differences based on the presence of children in the household. Mothers are less likely to have access to paid sick leave than women without children, whereas fathers are more likely to have access than men without children. In fact, the slight advantage in access for women seen in the aggregate is reversed such that fathers have more access than mothers. The fact that in 2010, 31% of mothers work part time compared to only 9% of fathers explains at least part of this reversal.

State-mandated paid sick leave is a gender-neutral policy, yet to the extent that men and women differ in their labor market decisions, these policies have the potential to differentially change access for men and women. Moreover, equal access need not be paired with equal take-up. Such differences provide the motivation for us to study gender differences in leave-taking for own illness and for care-giving following the implementation of state-mandated paid sick leave.

3. Data: Current Population Survey

Our analysis uses data from the 2006–2019 basic monthly files of the Current Population Survey (CPS). The CPS monthly files comprise a panel where each individual is interviewed once-per-month for four consecutive months, not interviewed for eight months, and then interviewed once-per-month again for four more consecutive months. The CPS basic monthly files contain detailed information about labor force participation, employment, usual weekly hours worked, and actual hours worked in the prior (reference) week. In addition, respondents are asked the reason for an absence from work during the reference week, where relevant. The basic monthly files also contain a host of demographic and geographic information about each individual, including gender, age, household composition, and links to other household members.

Our primary outcomes of interest are absences from work due to own sickness, or "sick leave," and absence due to care-giving. We identify absences using several questions related to a respondent's reported time at work. Certain absences are directly observable in the data using the survey question asking if the respondent was absent from work during the entire reference week. Workers who respond yes to that question are then asked the reason for their absence, which we use to differentiate between own illness and care-giving leaves. Specifically, we identify own sickness if respondents chose "Own illness" as the reason for their absence. We identify care-taking related absences as the result of "Child care problems" or "Other family/personal obligations."

To identify shorter leaves, we must infer an absence by comparing respondents' usual hours worked with their actual hours worked. We denote individuals who worked less than 35 hours during the reference week, but who typically work more than 35 hours, as taking leave. As with full week leaves, we can attribute leave to sick leave and care-giving leave because workers who worked less than 35 hours are asked the reason for working part time. For workers whose usual and actual hours are both less than 35, we identify a sick leave absence if they report their own sickness as the reason. Similarly, we identify a care-giving absence if they report either child care problems or other family/personal obligations as the reason for the absence. Finally, we combine full week and partial week absences (of the same type) to identify the full set of workers taking leave in the reference week.

All of our results are derived from the sample of adults aged 25 to 49 employed in the private sector.[7] To avoid the influence of seasonal patterns in leave-taking, our primary analysis is conducted at an annual level. We aggregate monthly absence data by identifying the set of workers who ever report an absence during the calendar year separately by type.[8] For all other variables, we keep the first monthly observation in a given year. Appendix Table A.2 provides a detailed description of all variables used in our analysis.

[7]We exclude federal, state, and local employees from our analysis because these employees are excluded from state paid sick leave mandates. Generally, these employees already have access to paid sick leave through their employer, regardless of state mandates.

[8]Individuals in the CPS are assigned inverse-probability weights for each month they are surveyed. When studying annual outcomes, we use the maximum monthly weight within the calendar year to re-weight our data to reflect the U.S. population.

Table 1 reports summary statistics for observations from 2006–2019 for those who live in states without a mandate and those living in a state with a mandate using only pre-treatment years. Column (1) reports statistics for our full analysis sample, and columns (2) and (3) report statistics for men and women, respectively. From column (1), we see that individuals are, on average, 36.4 years old and 56% are married. In addition, among those individuals with a spouse or cohabiting partner ("couples"), 77% of employed individuals are in dual-earning households. Among this population, 24% have preschool-aged children in their household, defined by the youngest child in their household aged five and younger, 28% have school-aged children in their household, defined by the youngest child in their household aged six through eighteen, and 49% do not have any children under the age of 18 in their household. In addition, 34% of our analysis sample has at least a bachelor's degree, 61% are white (non-Hispanic), 13% are black (non-Hispanic), and 19% are Hispanic. Finally, individuals in our sample are most likely to be employed in Professional, Management, Service and Office occupations (21%, 16%, 16%, and 12%, respectively).

Columns (2) and (3) allow us to compare the observable characteristics of employed men (53% of observations) and women (47% of observations) in our sample. Employed men, compared to employed women, are more likely to be married (57% compared to 53%) and, if they are in a household with a spouse or cohabiting partner, are more likely to live in a dual-earning household (88% compared to 69%). While there is no strong difference in the likelihood that employed men and employed women have preschool-aged kids (24% compared with 23%), employed men are less likely to have school-aged children (24% compared with 33%) and are less likely to have at least a bachelor's degree (32% compared with 37%). Employed men are also less likely to be black (11% compared with 15%) and more likely to be Hispanic than women (21% compared with 16%). Finally, men are less likely to be employed in Professional (17% compared with 25%), Service (12% compared with 20%), and Office occupations (6%

compared with 20%) and are more likely to be employed in Construction (11% compared with 0%) and Transportation (10% compared with 2%).

In our analysis sample, 89% of individuals are usually employed full-time as opposed to part-time. Men, however, are 12 percentage points more likely to be full-time employed than women. Turning to our outcomes of interest, 8% of employed individuals report having been absent for care-giving reasons in the prior week, and women are overwhelmingly more likely to take these absences (14% compared with 2%). Finally, 5% of individuals report having been absent due to their own illness during the prior week. Here again, women are slightly more likely to take this leave and the difference by gender is minimal (4% for men compared with 6% for women).

Figure 3 plots annual leave-taking rates from 2006–2019 separately for individuals in our analysis sample who live in states that passed a mandate by 2018 (solid line) and for those who live in states that have not passed a mandate by 2018 (dashed line). Panel (a) plots annual care-giving leave for our analysis sample, and panel (b) plots annual care-giving leave separately for men and women. Likewise, panel (c) plots annual sick leave rates for our analysis sample, and panel (d) separates these trends by gender. The vertical dashed line in these graphs marks the first time that a state (Connecticut) implements a paid-sick leave mandate in the U.S.

Panels (a) and (c) reveal that care-giving and sick leave taking rates are similar for states that do and do not pass a mandate for sick-leave and that these trends evolve similarly over time. Panels (b) and (d) further illustrate the stark differences in leave-taking by gender. In panel (b) we see that women are more than five times more likely to take leave for care-giving; by comparison, differences in sick leave-taking are much more muted by gender (panel d). In addition, these trends do not appear to be sharply different by gender for those individuals living in states that did and did not implement paid sick leave mandates, especially prior to 2012.

4. Two-Stage Difference-in-Differences Empirical Model

We identify the effect of mandated access to paid sick leave by comparing the leave-taking behavior of individuals based on whether or not they live in a state with a paid leave mandate around the timing of the policy change. This difference-in-differences comparison is typically carried out using a two-way fixed effects estimator, which permits an aggregation of evidence from separate quasi-natural experiments. In our setting, we combine evidence based on policy changes in 10 states and the District of Columbia between 2009 and 2018.

Typically, the average treatment effect based on a staggered treatment is estimated using a two-way fixed effects specification:

$$\text{Leave}_{ist} = \gamma_s + \lambda_t + \beta D_{st} + \varepsilon_{ist} \tag{1}$$

The variable Leave_{ist} is an indicator identifying one of two outcomes: that an individual i while living in state s during calendar year t reports being absent from work because of their own sickness or because of care-giving reasons. State fixed effects, γ_s, capture time-invariant differences across states, and year fixed effects, λ_t, account for overall trends in care-related absences. Finally, D_{st} is a treatment dummy that takes on a value of one for states in years when its state-mandated paid sick leave policy is in effect. In this model, $\hat{\beta}$ captures the average treatment effect.

Likewise, the dynamic treatment effect is captured using the following specification:

$$\text{Leave}_{ist} = \gamma_s + \lambda_t + \sum_{\tau=-q}^{-2} \alpha_\tau D_{ist} + \sum_{\tau=0}^{q} \beta_\tau D_{ist} \varepsilon_{ist} \tag{2}$$

Relative to the static model, the dynamic two-way fixed effect model provides empirical evidence of (1) pre-treatment differential trends across individuals in treated and control states up to q years before treatment based on the estimates of $\hat{\alpha}_\tau$, and (2) the dynamic profile of the treatment effect up to q periods after treatment based on the estimates of $\hat{\beta}_\tau$.

Several assumptions are required for the two-way fixed effect model to recover an unbiased and consistent estimate of the average treatment effect of access to paid sick leave. First, as is common for standard difference-in-differences models, identification requires an assumption of parallel trends. At the same time, because two-way fixed effects models combine many simple, or "2x2" difference-in-differences experiments, over time, additional assumptions about the nature of the treatment effect are also required (Goodman, 2021). Specifically, the treatment effect itself must be homogeneous across the sub-experiments; if the treatment effect differs by group, (e.g. because of differences in post-treatment dynamics or because the average treatment effect varies over time and the policy changes occur in different years), then this will induce bias in the traditional two-way fixed effects model.

To guard against these identification threats, our baseline estimates are based on the two-stage difference-in-differences estimator developed by (Gardner, 2022). As the name suggests, this estimator proceeds in two stages. In the first stage, we estimate the following specification:

$$\text{Leave}_{ist} = \gamma_s + \lambda_t + \varepsilon_{ist}$$

using only observations where $D_{ist} = 0$. In other words, we estimate the state and year fixed effects using observations from individuals who either (1) live in states that never implement state-mandated paid sick leave or (2) live in states that will eventually implement state-mandated paid in years before the mandate is in place. From these estimates, we residualize the data using all observations, including those for which $D_{ist} = 1$ using

$$\widehat{\text{Leave}}_{ist} = \text{Leave}_{ist} - \hat{\gamma}_s - \hat{\lambda}_t.$$

The second stage of the estimation compares residualized absences across treated and control units:

$$\widehat{\text{Leave}}_{ist} = \beta D_{st} + \varepsilon_{ist}$$

where $\hat{\beta}$ is an unbiased and consistent estimator of the average treatment effect of the likelihood of taking sick leave or care-giving leave for those who have access to paid sick leave mandates, and similarly,

$$\widehat{\text{Leave}}_{ist} = \sum_{\tau=-q}^{-2} \alpha_\tau D_{ist} + \sum_{\tau=0}^{q} \beta_\tau D_{ist} + \varepsilon_{ist}$$

yields unbiased and consistent estimators, $\hat{\beta}_\tau$, of the average treatment effect in each post-treatment period.

In additional analyses, we implement these same models to study the effect of access to paid sick-leave among several subsamples of our data including by gender, by race and ethnicity, by education, by martial status, and based on the relative earnings of adults within the same household.

5. Results

5.1. Effect of Mandated Paid Sick Leave: Leave Taking

Table 2 shows the estimated effects of mandated access to paid sick leave on work absences based on the two-stage differences-in-differences model. This model is our preferred specification because, as described in Section 4, it is robust to threats to identification driven by heterogeneous average treatment effects across groups and over time.[9] All estimates are based on employed adults in the private sector aged 25–49 between 2006 and 2019. Control means reflect the likelihood of care-taking leave measured using individuals in states that never implement a paid sick leave mandate and individuals in states that implement a paid sick leave mandate in the years prior to implementation. In all specifications we include controls for age, race, occupation, and marital status, and we report bootstrapped standard errors below estimates in parentheses.

Our estimates suggest that these mandates had little impact on the likelihood of taking either care-giving leave or sick leave across the full population. Panel A reports estimates for the likelihood that an employee takes leave for reasons related to care-giving, and panel B reports estimates for the likelihood that an employee takes leave for reasons related to own illness. Recall that work absences are a fairly uncommon event: 8.05% of employees in our analysis sample report having been absent from work for care-giving reasons and 4.76% report having been absent due to own illness (column 1, control mean). We find that the likelihood of care-giving leave increased by just 1% (0.0892 percentage points / 8.05%, panel A col 1), and this estimate is imprecisely measured (t-statistic of 0.236). In addition, we estimate that

[9]For the interested reader, Appendix Table A.3 reports estimates based on the two-way fixed effects estimator. This table reveals that our estimates similar across these two empirical models. Combined with the fact that staggered treatment happens over a short period, this suggests that threats to identification due to heterogeneous average treatment effects over time are not of first-order importance.

the likelihood of sick leave increased by 2% (0.105 percentage points / 4.76%, panel B col 1); this estimate is also statistically noisy, with a t-statistic of 0.656.

Next, we study the effect of paid sick leave mandates on work absences by gender. As previously described, men and women sort into different types of jobs, which can result in disparate access to paid sick leave through their employer. Therefore, men and women are likely to be differentially affected by these mandates. Moreover, gender norms in the United States are such that women are more likely to serve in a care-giving role, both for children and other family members, than men. Consistent with this, women are five times more likely to have been absent from work in the prior week for care-giving related reasons (2.32% for men and 14.5% for women, panel A). By comparison, there is less reason to expect that women and men contract illness at different rates. Consistent with this, we find similar baseline likelihoods of sick leave-taking for men and women: 3.89% of men are likely to report having been absent in the prior week due to own illness compared with 5.75% of women.

Columns 2 and 3 of Table 2 show that the small and statistically insignificant estimate for the full sample of adults masked considerable heterogeneity by gender. We find that men are 14% more likely to take leave for care-giving reasons (0.328 percentage points / 2.32%, panel A col 2), and this estimate is statistically significant at the 1% level (t-statistic 3.55). Next, we find that women are 2% *less* likely to take leave for care-giving reasons (-0.221 percentage points / 14.5%, panel A col 3), an economically small and statistically imprecise (t-statistic of 0.293) result. By comparison, we do not find evidence for a differential effect on taking sick leave by gender. For both men and women, similar to the full sample, we estimate that the likelihood of sick leave taking increased in the range of 1.5–2.5%, and these estimates are statistically insignificant at all conventional levels (panel B, cols 2 and 3).

Figure 4 plots event study coefficients describing the post-treatment dynamics of the effect of mandated access to paid sick leave for four years before implementation through four years after implementation.[10] In this analysis, all estimates reflect changes relative to the year before implementation, which is the reference period. This event study window reflects the fact that state paid sick leave policies are a more recent policy innovation, which limits our ability to study a lengthier post-treatment period. Panels (a) and (b) depict estimates of these effects on care-giving leave and sick leave respectively for our full analysis sample. These panels show that there are effect dynamics for our full analysis sample, consistent with our estimated null effect averaged across all post-treatment years. Panels (c) and (d) depict estimates for men and and panels (e) and (f) depict estimates for women. Corresponding point estimates are reported in columns 1, 2, and 6 of Appendix Table A.4 and Appendix Table A.6.[11]

Starting with care-giving leave, we estimate that the likelihood of leave-taking for care-giving by men increases from 13% one year after the mandate (0.283 percentage points / 2.24%, col 2) to 22% three years after the mandate (0.493 percentage points / 2.24%, col 2), and these estimates are statistically significant at the 5% level.[12] By comparison, we estimate that the likelihood of taking care-giving leave by women decreases by 4% (0.511 percentage points/13.6%, col 6) one year after the mandate and remains effectively unchanged from that estimate (0.00796 percentage point decrease, col 5) three years after the mandate passed; in all cases, these estimates for women are imprecisely measured. Based on this evidence, we conclude that care-giving leave increased after the mandate for men but not for women.

The event study estimates also allow us to test for parallel pre-trends in support of the identifying assumption of the difference-in-differences estimator. Specifically, we test for

[10]Our dynamic specification captures all pre and post treatment dynamics, including years outside of the -4 to 4 event time window. However, estimates for the dynamic treatment effect outside this window are only identified by the DC and CT expansion, which affected only a small share of the US workforce. For this reason, we suppress coefficients for treatment effects outside of the -4 to 4 window.

[11]Reported control means reflect average leave-taking for individuals in treated states in the year prior to treatment. Reported p-values reflect a parallel pre-trends test, or the null hypothesis that there was no difference in leave-taking across all years prior to treatment.

[12]We note that our estimates of the dynamic post-treatment effects are not statistically different from each other.

statistical evidence to reject the null hypothesis that the anticipation effect prior to treatment in the first four pre-treatment years is jointly zero. We fail to reject this null hypothesis at the 5% level for both men (p-value 0.411) and women (p-value 0.0721).

Panels (b), (d), and (e) of Figure 4 depict the post-treatment dynamics of the effect of the mandate on the take-up of sick leave. As in our static differences-in-differences estimate, our estimates of the post-treatment dynamic effect are small and statistically insignificant for both men and women. Importantly, these graphs expose the fact that there were trends in the take-up of paid sick leave even before the implementation of the mandate, in the full sample and in the separate subsamples of men and women. For sick leave take-up, p-values associated with the null hypothesis of parallel pre-trends for men and women are 0.00212 and 0.0000775 respectively. In other words, we can reject this null hypothesis at all conventional levels. For this reason, the estimated treatment effects for sick-leave take-up from our samples should be interpreted with caution.

5.2. Effect of Mandated Paid Sick Leave: Role of Children

We provide evidence that access to paid sick leave increases the likelihood that men are absent from work due to care-giving leave: men are 14% more likely to take care-giving leave after a mandate passes. In this section, we decompose our estimate based on the presence of children in the household. Specifically, we study whether there exists variation in the mandate's impact based on the age of the youngest child in the household. Accordingly, we create three mutually exclusive groups: those for whom the youngest child is preschool aged (≤ 5), those for whom the youngest child is in formal schooling (6–18), and those who do not have any children (18 or younger) in the house. We hypothesize that care-giving needs for those with children are likely to be greater than for those without children in the house. Table 3 reports estimated effects of the policy on the likelihood of care-giving leave separately for those who report the presence of preschool aged kids (cols 1, 2), the presence of school-aged children (cols 3, 4), and for those who do not have children in their household (cols 5, 6). All estimates are based on a two-stage differences-in-differences model and include controls for age, race, occupation, and marital status, following our baseline model.

Table 3 reveals that the strongest effect of the policy is concentrated among men with children. Men with preschool-aged children are 18.9% more likely to take care-giving leave (0.627 percentage points/3.32%, col 1) and men with school-aged children are 18.5% more likely to take care-giving leave (0.411 percentage points/2.22%, col 3) after the policy change. By comparison, men without children in the household experience an increase in care-giving leave that is roughly two-thirds as large as men with children (12%) and only marginally significant (t-statistic 1.90). In other words, our finding that men benefit from state-mandated paid sick-leave by increasing leave for care-giving is driven by men with children in their house. As in our baseline analysis, the separate estimates by presence of children are much noisier for women and not statistically different from zero for all groups.

Finally, Figure 5 plots the dynamic post-treatment effect of the mandates on care-giving leave for men (panels a–c) and women (panels d–f) splitting results by the presence of children. Corresponding point estimates are reported in Appendix Table A.4.[13] Panels (a) and (b) illustrate that state mandates lead to a persistent increase in the likelihood that men with children in their household take care-giving leave after the policy is effective. For men with preschool aged children, the effect of the mandate increases over time from a 12% increase

[13]For the interested reader, Appendix Table A.5 reports estimates of the effect of the mandate on sick leave take-up by gender according to whether or not there are children in the household In most cases, pre-trends are problematic, making these estimates difficult to interpret. One notable except is for men with school-aged children, who are 20% more likely to take sick leave (0.794/3.86%, col 3), an estimate that is statistically significant at the 1% level. If men with school aged children are taking care-giving leave to care for a sick child, this could increase the likelihood that they themselves contract an illness, which in turn could increase the likelihood that they take sick leave.

(0.437 percentage points/3.5%, col 3) one year after the mandate passed to a 30% increase (1.04 percentage points/3.5%, col 3) three years after the mandate passed. For men with school aged children, the effect of the mandate was both large and stable throughout the post-treatment period, with a roughly 40% increase in the likelihood of taking care-giving leave two years after the mandate (0.779 percentage points/2.04%, col 4) that remains of similar magnitude in years three and four.

5.3. Effect of the Mandated Paid Sick Leave: Variation by Race, Education, and Other Household Characteristics

We have shown that men are 14% more likely to take care-giving leave after a mandate passes, and that this result is driven by a 19% increase among men with children under 18 in their household. In this section, we study variation in the effect of treatment driven by demographic differences across individuals that could correspond to inequities in access to paid sick leave prior to the implementation of state mandates. This analysis exposes additional factors interacting with the policy change that may affect the likelihood of care-giving leave.[14]

Figure 6 plots trends in *access* to paid sick leave for men and women based on race and ethnicity (panels a and b), education (panels c and d), and household structure (panels e and f) based on NHIS data. As before, the dashed line in 2012 marks the first state mandated paid sick leave policy (Connecticut). Figure 7 reports the estimated effect of state mandates on care-giving leave based on these same demographic variables. In this figure, we report scaled estimates, which correspond to the estimated effect of the mandate scaled by the average take-up of care-giving leave for untreated individuals and treated individuals in years prior to the mandate.

[14] We focus this analysis on care-giving leave rather than sick leave in light of our previous analysis, which exposed problematic pre-trends in almost all analysis.

Beginning with the racial and ethnic decomposition, panels (a) and (b) highlight that substantial differences in access to paid sick leave existed for Hispanic workers (dotted line) as compared to White and Black workers (solid and dashed lines, respectively). Before 2012, just 40% of Hispanic men had access to paid sick leave, compared with roughly 60% of White and Black men, and 50% of Hispanic women had access compared with 60% of White and Black women. Notably, these gaps shrink dramatically after 2014, which coincides with the expansion of state mandates. These patterns are consistent with Bartel et al. (2023), who find that Hispanic workers have lower rates of paid family and medical leave access and use than White, Non-Hispanic workers.

Consistent with this, panel (a) of Figure 7 reveals the largest estimated effect of the mandates is for Hispanic men. Hispanic men are 25% more likely to take care-giving leave after a mandate is passed, and this estimate is statistically significant at the 1% level. We also estimate a small, positive effect for Hispanic women (3%); while this estimate is not statistically significant, it stands out as the only positive effect for women when results are split by race and ethnicity.

Next, we study whether variation exists based on an individual's educational attainment. Panels (c) and (d) of Figure 6 plot the share men and women respectively with access to paid sick leave, splitting by those with less than a bachelor's degree (solid line) and those with at least a bachelor's degree (dashed line). Prior to the implementation of any state mandates, men with less than a bachelor's degree were substantially less likely to have access to paid sick leave (50% compared with 80%). Similar disparities exist for women, although the magnitude of these gaps is smaller. After 2012, however, the gap in paid sick leave access appears to shrink for men. Accordingly, panel (a) of Figure 7 reports that men with less than a bachelor's degree are 18% more likely to take care-giving leave after a mandate is passed, and this estimate is statistically significant at the 1% level. By comparison, we estimate that men with a bachelor's degree are just 11% more likely to take care-giving leave, and this estimate is statistically insignificant at all conventional levels. We find no meaningful differences in the likelihood of taking care-giving leave among women in the post-treatment period, irrespective of educational attainment (Figure 7, panel b).

Finally, panels (e) and (f) of Figure 6 plot access to paid sick leave for men and women respectively based on variation in dual earner status. We categorize individuals as "Single" if they do not report a spouse or a cohabiting partner in the same household (solid line). For individuals with a spouse or a cohabiting partner, we group them by whether they are the only one in the household working (single earner, dashed line) or by whether both partners are working (dual earner, solid line). Single men and single earning men are less likely to have access to paid sick than dual-earning men leave prior to 2012. This pre-exisiting disparity in access is consistent with findings in panel (a) of Figure 7, which reveal that both single men and single-earning men drive our estimate. We find that single men are 17% more likely to take care-giving leave, and that single-earning men are 19% more likely to take care-giving leave after the mandate is passed; each of these estimates is statistically significant at the 5% level. The descriptive relationships in access to paid sick leave are more variable among women, but single-earning women are the least likely to have access to paid sick leave. We estimate a that single-earning women are 4% more likely to take care-giving leave, but these results are noisy making it difficult to draw any strong conclusions.

In Appendix Table A.6, we find evidence that men with children in their household are 15% more likely to take sick leave (cols 3 and 5). These results are consistent with increased care-giving leave to care for sick children, which would increase the likelihood of the care-giver themselves contracting an illness. On the other hand, we find weak statistical evidence that men without school-aged children in their household are 7% less likely to take sick-leave (t-statistic 1.93). As in all other estimates, we find no effect of mandated access to paid sick leave on sick leave take-up for women, regardless of the presence of children in their household.

5.4. Discussion of Results: Gender Gaps

In light of overwhelming evidence that men are more likely to take leave for care-giving reasons after state mandates are imposed, we ask whether these changes led to a reduction in the gender gap for care-giving leave. In appendix Table A.7 we report estimates of the effect of mandates on the gender gap based on a modification of Equation (1) that includes a dummy variable identifying females, and all of the relevant interactions using a two-stage difference-in-differences model. Estimates are reported for the full sample in column (1) and the separately decomposed based on the presence of young children in columns (2) – (4). Although these estimates are statistically imprecise, likely due to to the high variance in the effect of the mandate for women, the magnitude of these estimates suggests that the gap in care-giving between men and women shrunk by 5% (-0.00596 percentage points/12.2%, col 1) for the full sample, and by 4-5% among men with children (cols 2 and 3).

This result is especially notable given that women are penalized in the labor market for having children both compared to women without children Yu and Hara (2021) and relative to men with children Kleven et al. (2019). In the United States, gender norms are offered as one mechanism for this child penalty Kleven (2022). Importantly, while gender norms in the U.S. dictate that women are the primary care-takers of young children, increasing access to paid sick leave appears to have nudged some men into shouldering more of those efforts. From a policy standpoint, our results suggest that mandated paid sick leave may help equalize care duties across parents, which could in turn provide positive labor market returns for women who currently bear the brunt of care-giving.

6. Conclusion

This paper is one of the first to examine the effects of newly enacted state-level paid sick leave mandates on leave-taking behaviors for workers living in a broad set of states. In general, paid sick leave mandates guarantee that workers have access to paid leave benefits to be used for work absences due to own illness or caring for a sick family member. We find that on the whole, increasing access did not materially change the likelihood workers are absent from work due to their own illness or a family member's illness. However, we find substantial differences when we look at these effects by gender.

Our empirical analysis shows that despite being a gender-neutral policy, mandated access to paid sick leave has gendered implications for leave-taking. Overall, paid sick leave mandates increase the probability of care-giving leave by men, and more specifically for men with children and men who were ex-ante less likely to have had access to paid sick leave (Hispanic workers and men without a bachelor's degree). By contrast, women are no more likely to use leave for care-giving after the policy change. While paid sick leave policies are typically motivated from a public health standpoint to minimize contagion of illness, the gender differences we find for care-giving leave suggest a possible auxiliary benefit. In particular, recent literature suggests women bear a labor market cost to having children. To the extent that women are penalized through self-selection into family-friendly jobs with leave benefits but lower wages, an equalization of leave benefits could help ease the burden of care-giving and possibly reduce wage pressures faced by women. We leave it to future work to investigate the impact of paid sick leave mandates on the gender-wage gap.

Our results contribute to a nascent and growing literature studying the effect of paid sick leave mandates. For example, Stearns and White (2018) find that the mandates in Connecticut and the District of Columbia decreased leave taking for own illness. The authors attribute this result to a decline in the transmission of illness due to increased leave-taking, which, on net, led to an aggregate reduction in sick leave take-up. Maclean, Pichler and Ziebarth (2020) study the impact of state sick leave mandates on access, use, and labor costs using restricted-access, job-level data contained in the National Compensation Survey; these data show that newly covered employees took two additional leave days per year at little additional cost to the employer. Finally, Byker, Smith and Patel (2022) study the impact of the emergency paid sick leave that was temporarily mandated during the onset of the pandemic, finding that mandated access allowed workers to accommodate sudden increased need for sick leave. We complement and expand this body of work by studying the differential impact of these mandates on men and women separately for sick leave and care-giving leave. Our findings provide important new context on areas of existing need for paid sick leave and the importance of such leave to accommodate short-duration work absences likely related to childcare.

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7. Tables and Figures

Table 1
Summary Statistics

	Analysis Sample (1)	Men (2)	Women (3)
Age	36.4	36.3	36.5
Married	0.56	0.57	0.53
Couple, Dual Earner	0.77	0.69	0.88
Age of Youngest Child			
0-5 years old	0.24	0.24	0.23
6-18 years old	0.28	0.24	0.33
No Children Under 18	0.49	0.52	0.45
Education			
At Least Bachelors Degree	0.34	0.32	0.37
Race			
White, Not Hispanic	0.61	0.60	0.61
Black, Not Hispanic	0.13	0.11	0.15
Hispanic	0.19	0.21	0.16
Other	0.08	0.08	0.08
Occupation			
Manangement	0.16	0.15	0.16
Professional	0.21	0.17	0.25
Service	0.16	0.12	0.20
Sales	0.11	0.10	0.11
Office	0.12	0.06	0.20
Farming	0.01	0.01	0.00
Construction	0.06	0.11	0.00
Maintenance	0.04	0.07	0.00
Production	0.07	0.10	0.04
Transportation	0.07	0.10	0.02
Employed Full Time	0.89	0.94	0.82
Work Absences			
Care-Giving Leave	0.08	0.02	0.14
Sick Leave	0.05	0.04	0.06
Observations	1,310,259	695,692	614,567

Notes: This table reports summary statistics for our primary analysis sample, which is comprised of private sector, employed individuals aged 25–49. Summary statistics are based on control observations from (1) states without mandated paid sick-leave from 2006–2019 and (2) pre-treatment years from states with mandated paid sick-leave. See Section 3 for more details about this analysis sample. Source: Current Population Survey

Table 2
Effect of Mandated Paid Sick Leave on Work Absences

	Full Sample (1)	Men (2)	Women (3)
Panel A: Care-Giving Leave			
Treatment Effect	0.000892	0.00328***	-0.00221
	(0.00378)	(0.000923)	(0.00755)
Control Mean	0.0805	0.0232	0.145
Panel B: Sick Leave			
Treatment Effect	0.00105	0.000684	0.00142
	(0.00160)	(0.00153)	(0.00194)
Control Mean	0.0476	0.0389	0.0575
Observations	1,399,581	742,375	657,206

Notes: This table reports estimates of the effect of access to mandated paid sick-leave on the likelihood of an individual being absent from work for either care-giving (panel A) or own-illness (panel B) reasons based on a two-stage differences-in-differences model. Treatment effect reflects $\hat{\beta}$ as described in equation 1. Control means measure average leave-taking among untreated observations. Estimates are reported for the full sample of employed adults aged 25–49, and separately for men and women. All specifications control for state and year fixed effects in addition to age, age^2, race, occupation, and marital status. Bootstrapped standard errors are reported based on 500 replications. *,**, *** indicate significance at 0.1, 0.05, and 0.01, respectively.

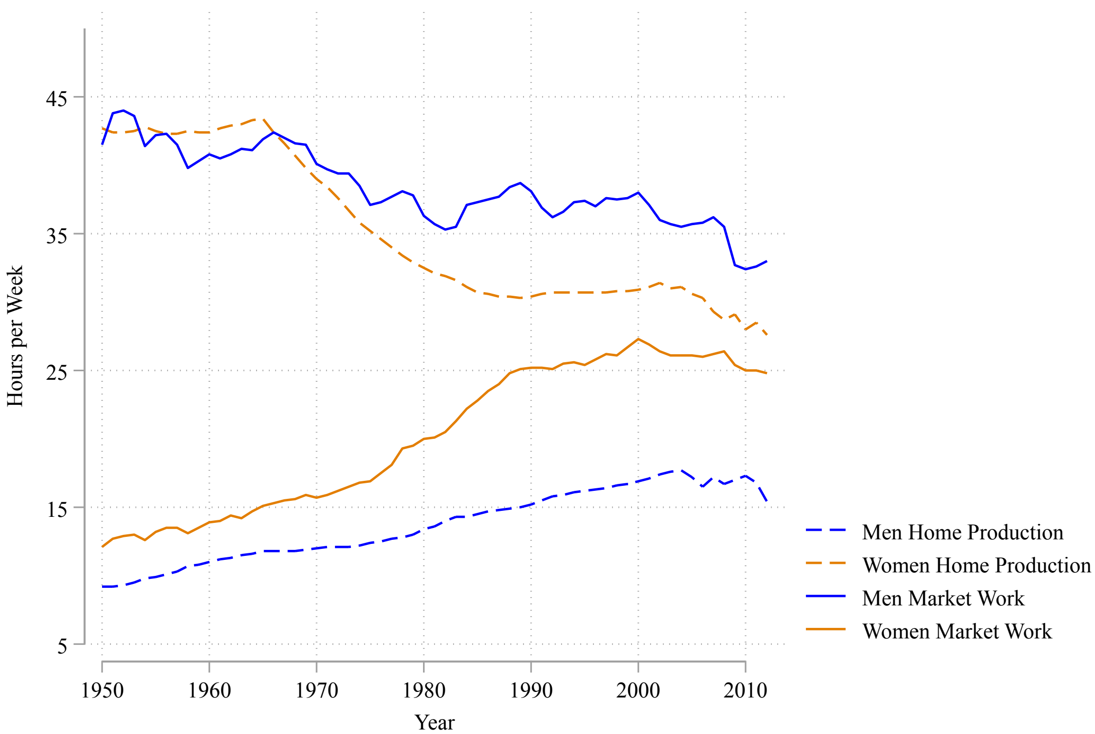
Table 3

Effect of Mandated Paid Sick Leave on Care-Giving Leave: Role of Children

	Youngest Child Aged ≤ 5		Youngest Child Aged 6–18		No Children Aged ≤ 18	
	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)
Treatment Effect	0.00627***	-0.00321	0.00411***	-0.00322	0.00224*	0.00400
	(0.00185)	(0.0109)	(0.00137)	(0.0121)	(0.00118)	(0.00314)
Control Mean	0.0332	0.256	0.0222	0.180	0.0189	0.0622
Observations	179,165	146,331	179,654	217,010	383,556	293,865

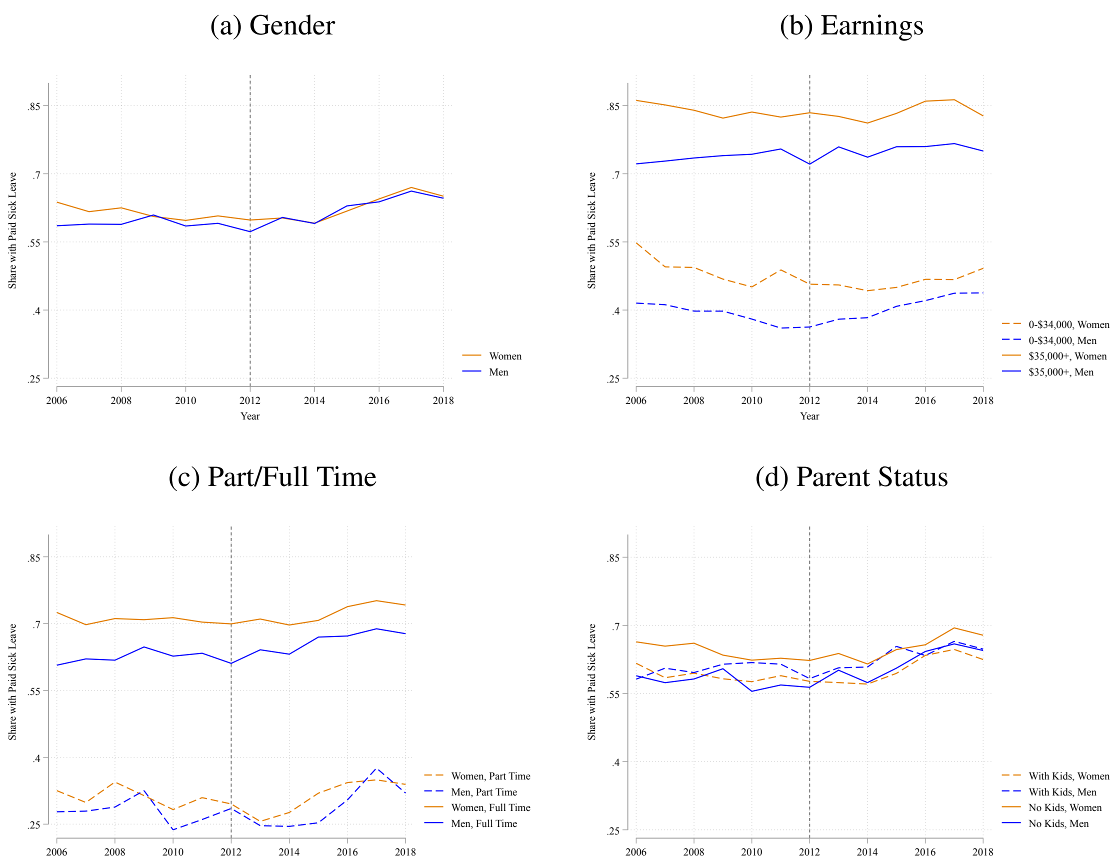
Notes: This table reports estimates of the effect of access to mandated paid sick-leave on the likelihood of an individual being absent from work for care-giving reasons based on the two-stage differences-in-differences model. Estimates are reported separately based on the age of the youngest child in the household. See also Table 2 Notes.

Figure 1. Historical Time Use, by Gender



Source: The data for this figure come from Valarie A. Ramey (https://econweb.ucsd.edu/ vramey/research.html#time) and provide an update to Ramey and Francis (2009) and Ramey (2009).

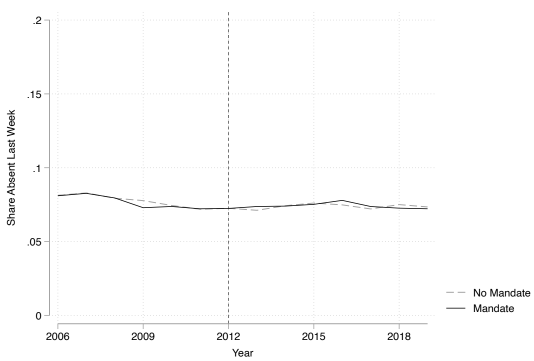
Figure 2. NHIS Sample: Trends in Access to Paid Sick Leave, 2006–2018



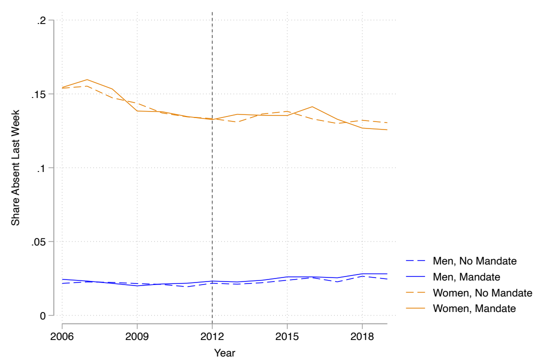
Notes:
Analysis sample based on civilian, employed adults aged 25–49. Paid Sick leave is a dummy variable equal to one for an individual reporting they "have paid sick leave at their main job" amongst those with a job last week. Source: National Health Interview Surveys.

Figure 3. CPS Analysis Sample: Work Absences Trends, 2006 – 2019

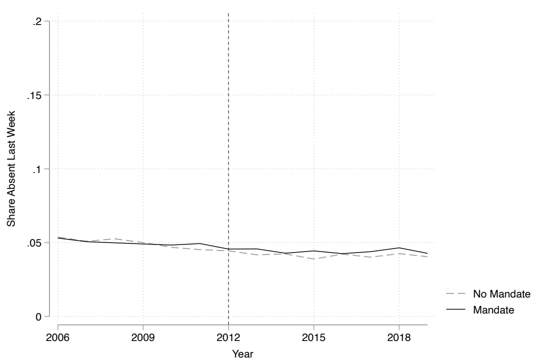
(a) Care-Giving Leave



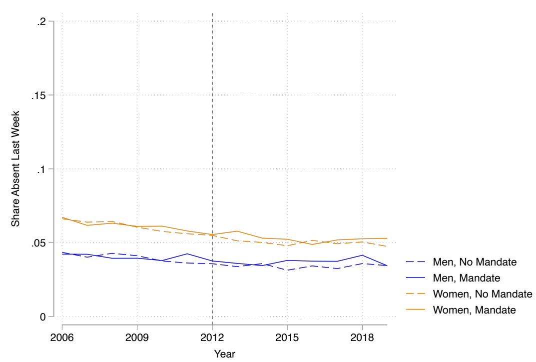
(b) Care-Giving Leave: By Gender



(c) Sick Leave

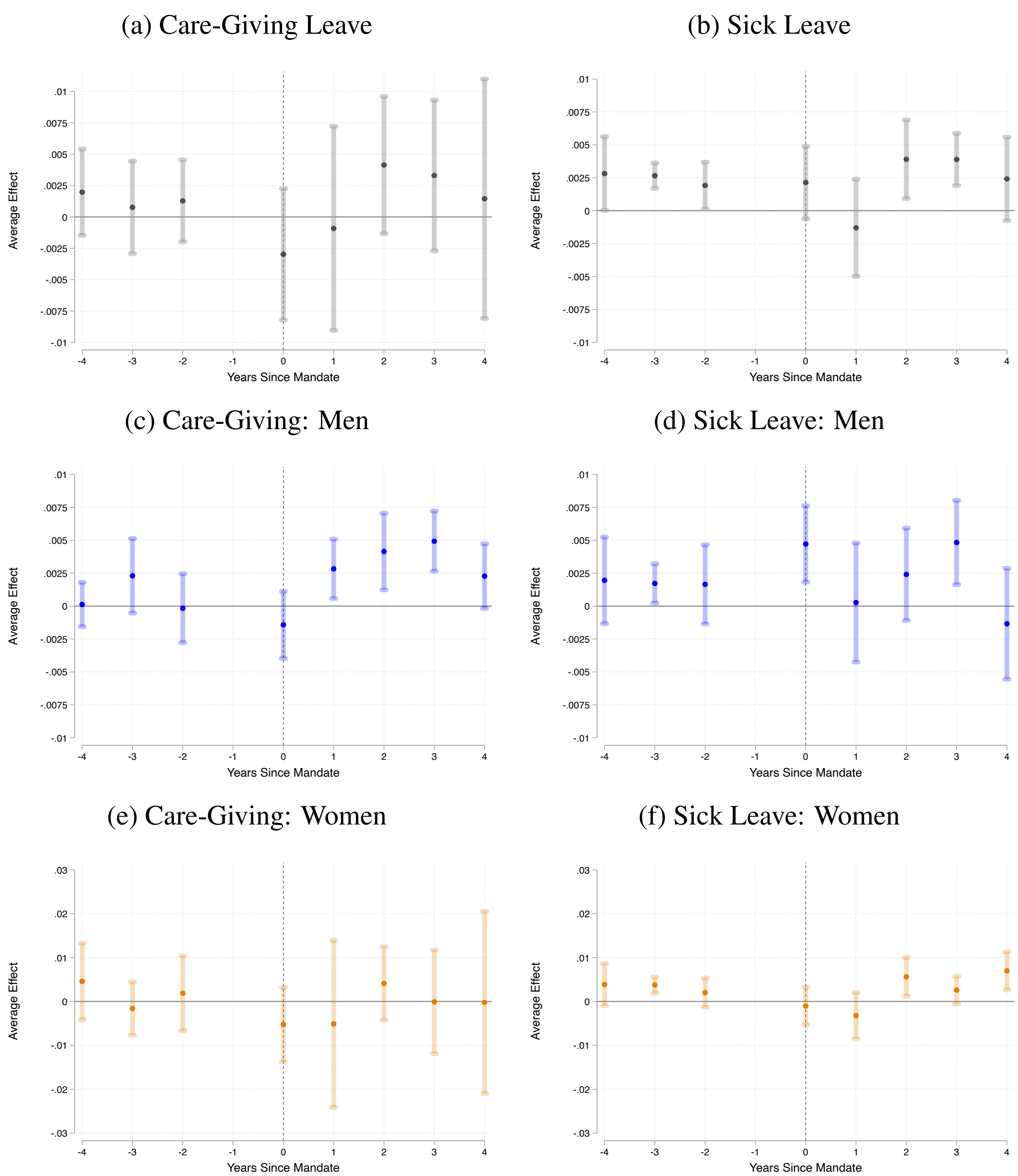


(d) Sick Leave: By Gender



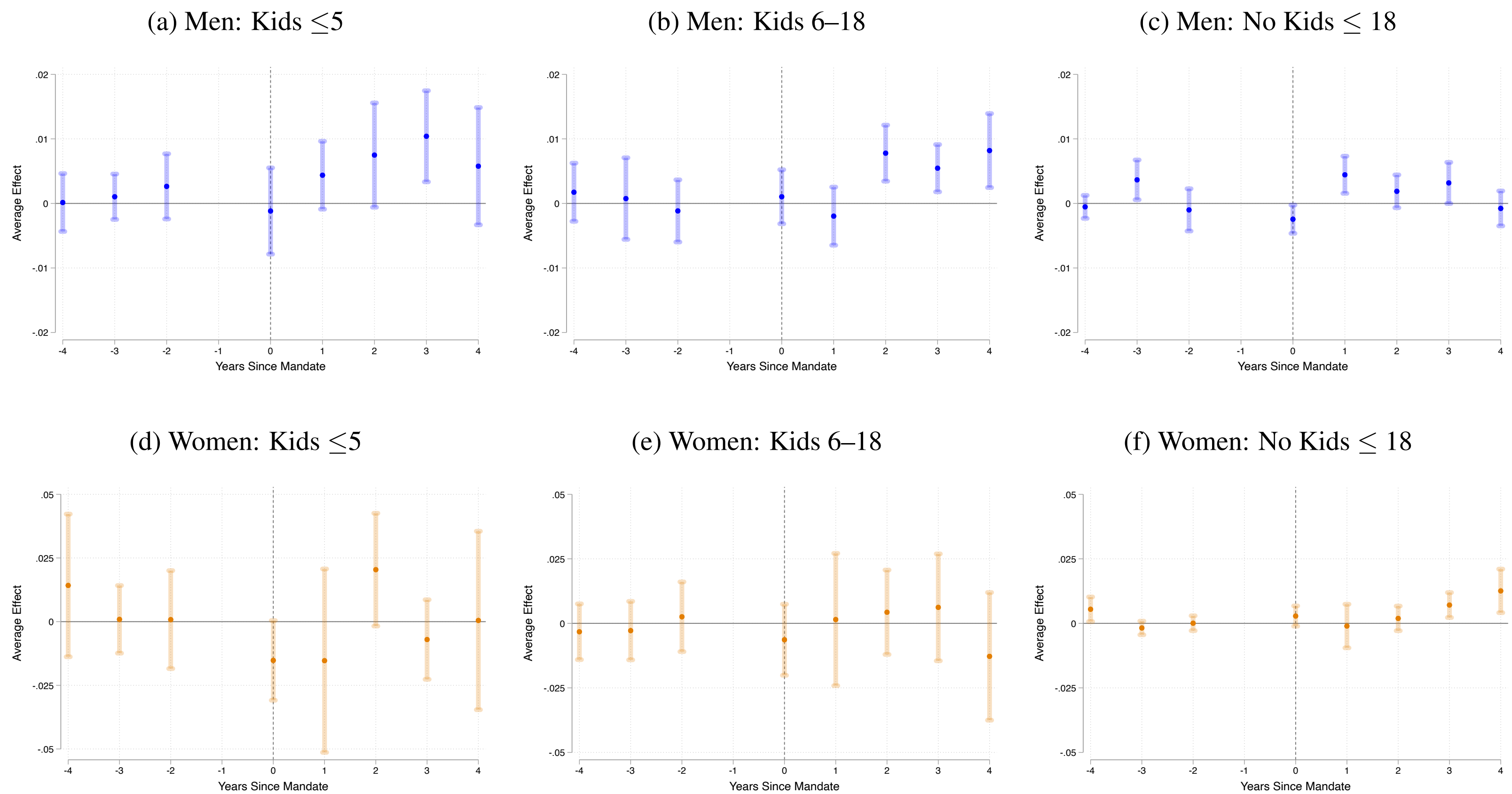
Notes: This figure plots the share of workers in the analysis sample who report having been absent from work in the prior week for care-giving reasons (panels a and b) and own-illness (panels c and d). Panels (c) and (d) plot leave-taking rates separately for men (triangles) and women (diamonds). Analysis sample is limited to employed adults aged 25–49; see Section 3 for details. Source: Current Population Survey.

Figure 4. Dynamic Post-Treatment Effect of State Mandates on Care-Giving Leave: Variation by Gender



Notes: This figure plots the dynamic treatment effect of mandated access to paid sick leave based on the dynamic two-stage differences-in-differences model described in equation 2. Estimates are reported separately for care-giving and sick leave and based on the gender of the employee. 95% confidence intervals are depicted by the shaded region. Point estimates are reported in Appendix Table A.4 and Appendix Table A.6. See Table notes for additional details. Source: Current Population Survey

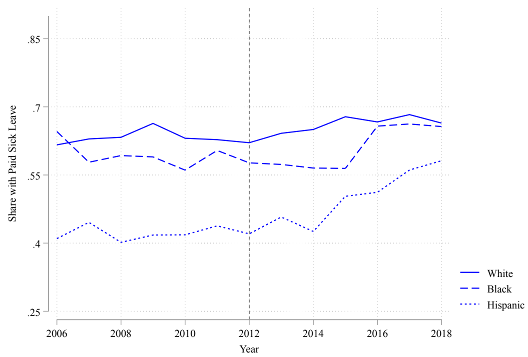
Figure 5. Dynamic Post-Treatment Effect of State Mandates on Care-Giving Leave: Role of Children



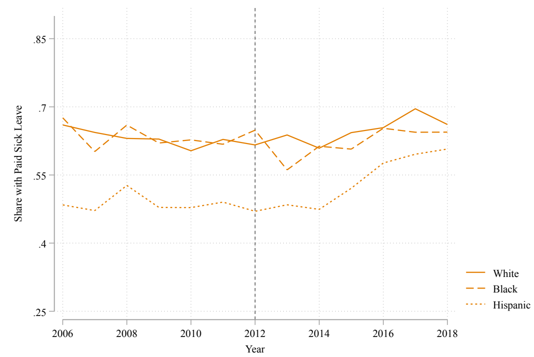
Notes: This figure plots the dynamic treatment effect of mandated access to paid sick leave based on the dynamic two-stage differences-in-differences model described in equation 2. All estimates are relative to the year before the mandate passed. Estimates are reported for men and women based on the age of the youngest child in their household. 95% confidence intervals are depicted by the shaded region. Point estimates are reported in Appendix Table A.4. See Table notes for additional details. Source: Current Population Survey

Figure 6. NHIS: Subgroup Trends in Access to Paid Sick Leave, 2006–2018

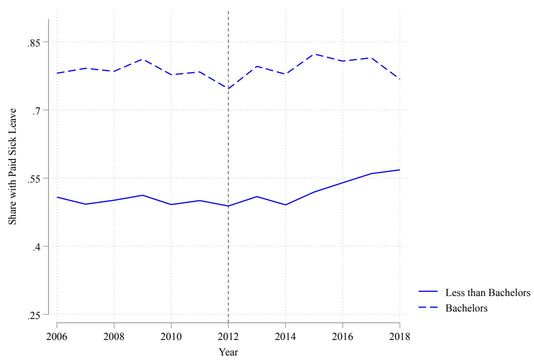
(a) Race and Ethnicity: Men



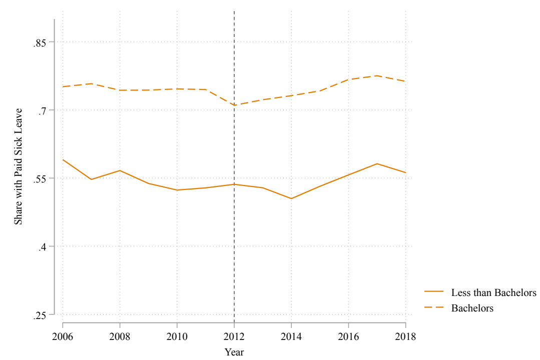
(b) Race and Ethnicity: Men



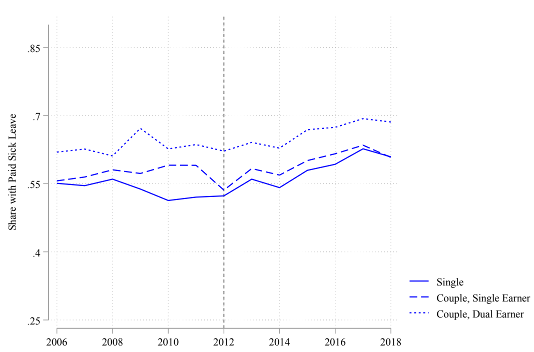
(c) Education: Men



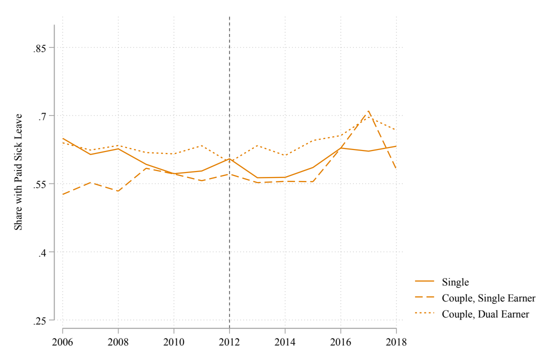
(d) Education: Women



(e) HH Status: Men



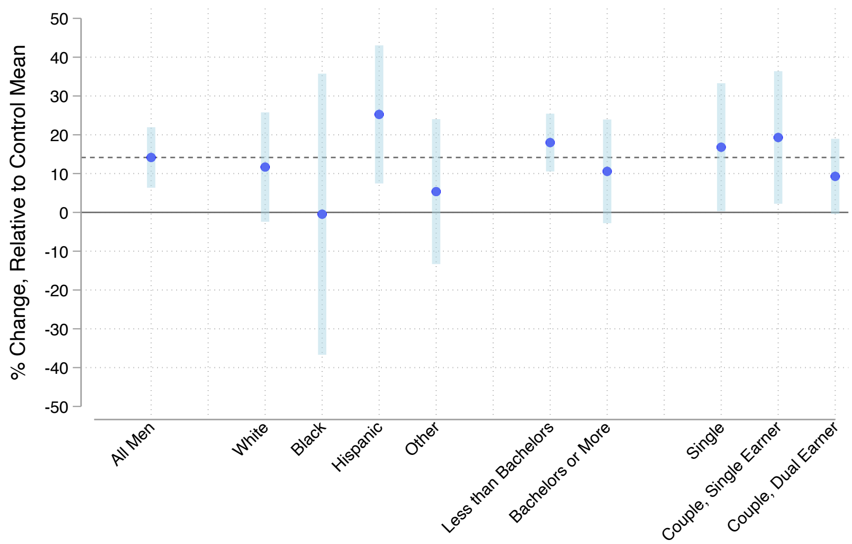
(f) HH Status: Women



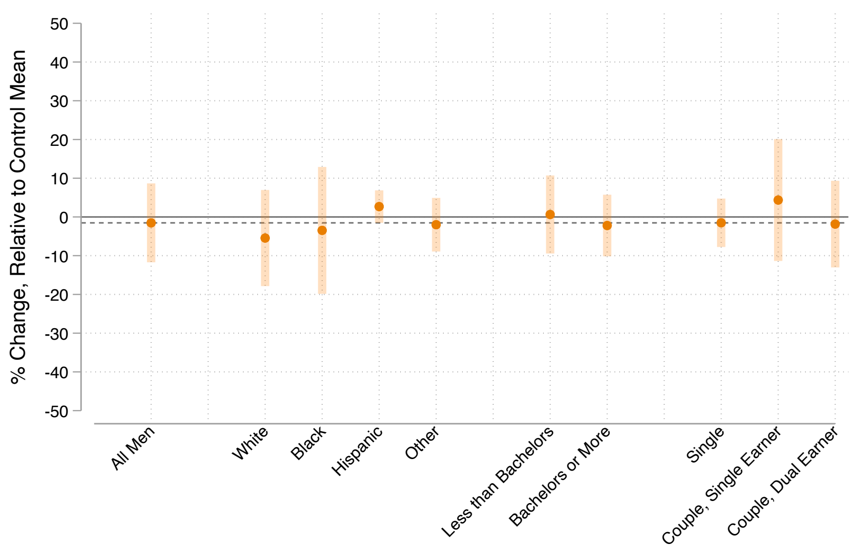
Notes:
Analysis sample based on civilian, employed adults aged 25–49. Paid Sick leave is a dummy variable equal to one for an individual reporting they ”have paid sick leave at their main job” amongst those with a job last week. Source: National Health Interview Surveys.

Figure 7. Post-Treatment Effect of State Mandates on Care-Giving: Subsample Aanlysis

(a) Men



(b) Women



Notes: This figure plots the average treatment effect of mandated access to paid sick leave based on the dynamic two-stage differences-in-differences model described in equation 2 across subsamples based on race, education, martial status, and intrahousehold earnings behavior. Estimates are scaled based on the control mean. 95% confidence intervals are depicted by the shaded region. See Table A.2 for additional details on subgroup defintitions. Source: Current Population Survey

Table A.1: State Paid Sick Leave Mandates

State	Covered Employers	Covered Employees	Year Implemented	Number of Employees Eligible	Frac of Employees Eligible (In State)	Frac of Employees Eligible (In US)
Washington D.C.	All	All	2008	356,330	1.000	0.003
Connecticut	Employers with 50 or more employees	Service workers who work at businesses with at least 50 employees	2012	1,259,650	0.521	0.012
California	Employers with 1 or more employees who work more than 30 days in a year in California.	Employees who work 30 or more days per year in California for the same employer	2015	13,455,918	1.000	0.116
Massachusetts	All employers. Local governments and those with fewer than 11 employees may provide unpaid leave.	All employees	2015	2,789,769	0.853	0.024
Oregon	Employers with 10+ employees (unpaid if fewer than 10) or employers in large cities (500,000+ population) with 6+ employees anywhere in the state (unpaid if fewer than 6)	All employees of companies with 10+ employees (unpaid if fewer than 10) or employers in large cities (500,000+ population) with 6+ employees anywhere in the state (unpaid if fewer than 6)	2016	1,360,784	0.820	0.011
Arizona	All	All	2017	2,248,266	1.000	0.019
Maryland	Employers with 15+ employees except if covered by local ordinance	Employees who work at least 12 hours per week	2018	1,966,677	0.729	0.016
New Jersey	All employers with workers in the state	All employees working for an employer in the state	2018	3,229,834	1.000	0.026
Rhode Island	Employers with 18+ employees in Rhode Island	All employees whose primary place of employment is in Rhode Island	2018	339,685	0.702	0.003
Vermont	All employers doing business or operating in the state	Employees who work for an average of no less than 18 hours per week during a year	2018	204,761	1.000	0.002
Washington	All	All	2018	2,383,137	1.000	0.019
Michigan	Employers with 50 or more employees	Employees who work at least 25 hours per week, who work at least 26 weeks per year for a job scheduled for at least 26 weeks, and whose primary work location is in Michigan	2019	3,452,817	0.604	0.028
Nevada	Employers in business for at least 2 years, with 50+ employees in the state	Employees of businesses with 50+ employees in the state	2020	1,159,738	0.568	0.009

Table A.1 – continued from previous page

State	Covered Employers	Covered Employees	Year Implemented	Number of Employees Eligible	Frac of Employees Eligible (In State)	Frac of Employees Eligible (In US)
New York	Private sector employers with 5+ employees or net income of more than $1 million	All employees of private sector employers with 5+ employees or net income of more than $1 million	2020	7,498,829	0.935	0.060
Colorado	Employers with at least 16 employees	Employees who work at businesses with at least 16 employees	2021	2,009,386	0.684	0.017
Maine	Employers with 10+ employees	All employees who work for covered employers accrue leave but ineligible to take it until 120 consecutive days of employment	2021	412,049	0.768	0.003
New Mexico	Private employer with at least 1 employee	All employees of private employers	2022	523,320	1.000	0.004
Illinois	All	All	2024	4,727,536	1.000	0.043

Notes: Sources: Paycor and Quarterly Census of Employment and Wages (QCEW). This table estimates the potential number of employees impacted by state sick leave mandates. The number of employees eligible in a state is estimated with Q1 employment level data aggregated by NAICS sector codes and firm size for a particular state and implementation year. Firms which do not meet either size or industry eligibility criteria as described in Columns 2 and 3 are excluded. The fraction of employees eligible in the state and the U.S. is then calculated by taking the number of remaining employees over the total state and national employment level as reported by the QCEW, respectively. This data does not include individual firm or employee characteristics that might impact eligibility, such as firm net income, employee state of residence, or average weekly hours worked. Illinois eligibility data is estimated using the latest available survey data from 2022.

Table A.2
Variable Definitions: Current Population Survey

Variable	**Definition**
Employed	Dummy variable equal to 1 for those individuals who were employed or had a job but not at work during the reference week. *Source: Basic Monthly Survey*
Part-Week Absence	Dummy variable equal to 1 for those individuals who report that they worked less than 35 hours during the reference week and also report that they typically work 40 hours per week. *Source: Basic Monthly Survey*
Full-Week Absence	Dummy variable equal to 1 for those individuals who report that they were absent from work during the reference week *Source: Basic Monthly Survey*
Ever Absent	Dummy variable equal to one if an individual ever reports a work-absence, either part-week or full-week, within a calendar year. *Source: Basic Monthly Survey*
Absent Due to Own Sickness	Dummy variable equal to one if an individual reports “Own Illness/Injury/Medical Problems” as the reason for their work absence during the reference week or “Own Illness” as the reason for working less than their usual hours during the reference week. *Source: Basic Monthly Survey*
Ever Absent, Own Sickness	Dummy variable equal to one if an individual ever reports that they are absent due to own sickness within a calendar year. *Source: Basic Monthly Survey*
Care-Related Absence	Dummy variable equal to one if an individual reports either “Child Care Problems,” “Other Family/Personal Obligations,” as the reason for their work absence during the reference week or as the reason for working less than their usual hours during the reference week. *Source: Basic Monthly Survey*
Ever Absent, Care-Related	Dummy variable equal to one if an individual ever reports a care-related absence within a calendar year. *Source: Basic Monthly Survey*
Age of Youngest Child	Age of youngest child under 18 in the household *Source: Basic Monthly Survey*
Occupation	Categorical variable sorting individuals into the following categories based on their reported occupation: Management, Professional, Service, Sales, Office, Farming, Construction, Maintenance, Production, Transportation, and Armed Forces. *Source: Basic Monthly Survey*
Race	Categorical variable sorting individuals into the following categories based on their reported race: White, Non-Hispanic, Black, Non-Hispanic, Hispanic, and Other. *Source: Basic Monthly Survey*
Bachelors Degree	Binary variable identifying individuals with at least a Bachelor’s Degree. *Source: Basic Monthly Survey*
Married	Binary variable identifying married individuals, whether their spouse is present or absent. *Source: Basic Monthly Survey*
Single	Binary variable identifying individuals who do not have a spouse or co-habitating partner present in the household. *Source: Basic Monthly Survey*
Couple, Single Earner	Binary variable identifying individuals who have a spouse or co-habitating partner present in the household who is not employed. *Source: Basic Monthly Survey*
Couple, Dual Earner	Binary variable identifying individuals who have a spouse or co-habitating partner present in the household who is employed. *Source: Basic Monthly Survey*

Table A.3
Effect of Mandated Paid Sick Leave on Work Absences: Two-Way Fixed Effects Model

	Full Sample (1)	Men (2)	Women (3)
Panel A: Care-Giving Leave			
Treatment Effect	0.000942	0.00336***	-0.00222
	(0.00373)	(0.000897)	(0.00752)
Control Mean	0.0805	0.0232	0.145
Panel B: Sick Leave			
Treatment Effect	0.00135	0.00110	0.00156
	(0.00142)	(0.00142)	(0.00168)
Control Mean	0.0476	0.0389	0.0575
Observations	1,399,581	742,375	657,206

Notes: This table reports estimates of the effect of access to mandated paid sick-leave on the likelihood of an individual being absent from work for either care-giving (panel A) or own-illness (panel B) reasons based on a two-way fixed effects model. See also Table 2 Notes.

Table A.4
Effect of Mandated Paid Sick Leave on Care-Giving Work Absences: Dynamic Estimates

		Men				Women			
	Full Sample (1)	Full Sample (2)	Youngest Child Aged ≤ 5 (3)	Youngest Child Aged 6–18 (4)	No Children Aged ≤ 18 (5)	Full Sample (6)	Youngest Child Aged ≤ 5 (7)	Youngest Child Aged 6–18 (8)	No Children Aged ≤ 18 (9)
4 Years Before Mandate	0.00197	0.000119	0.000142	0.00174	-0.000534	0.00457	0.0142	-0.00325	0.00546**
	(0.00176)	(0.000859)	(0.00229)	(0.00230)	(0.000910)	(0.00445)	(0.0143)	(0.00552)	(0.00244)
3 Years Before Mandate	0.000765	0.00230	0.00104	0.000752	0.00365**	-0.00161	0.000903	-0.00281	-0.00178
	(0.00188)	(0.00144)	(0.00180)	(0.00323)	(0.00157)	(0.00309)	(0.00678)	(0.00578)	(0.00135)
2 Years Before Mandate	0.00128	-0.000165	0.00264	-0.00116	-0.00100	0.00186	0.000804	0.00255	0.0000559
	(0.00167)	(0.00133)	(0.00258)	(0.00246)	(0.00167)	(0.00435)	(0.00982)	(0.00690)	(0.00147)
Year of Mandate	-0.00298	-0.00142	-0.00118	0.00104	-0.00245**	-0.00525	-0.0152*	-0.00635	0.00283
	(0.00268)	(0.00130)	(0.00341)	(0.00214)	(0.00112)	(0.00434)	(0.00799)	(0.00703)	(0.00200)
1 Year After Mandate	-0.000914	0.00283**	0.00437	-0.00197	0.00444***	-0.00511	-0.0153	0.00148	-0.00104
	(0.00415)	(0.00115)	(0.00269)	(0.00230)	(0.00147)	(0.00970)	(0.0184)	(0.0131)	(0.00431)
2 Years After Mandate	0.00414	0.00414***	0.00749*	0.00779***	0.00188	0.00411	0.0204*	0.00431	0.00192
	(0.00279)	(0.00149)	(0.00412)	(0.00222)	(0.00130)	(0.00427)	(0.0113)	(0.00835)	(0.00244)
3 Years After Mandate	0.00330	0.00493***	0.0104***	0.00546***	0.00318*	-0.0000796	-0.00702	0.00621	0.00710***
	(0.00306)	(0.00117)	(0.00360)	(0.00187)	(0.00163)	(0.00603)	(0.00797)	(0.0106)	(0.00247)
4 Years After Mandate	0.00145	0.00227*	0.00577	0.00819***	-0.000776	-0.000224	0.000469	-0.0128	0.0126***
	(0.00487)	(0.00125)	(0.00464)	(0.00293)	(0.00138)	(0.0106)	(0.0179)	(0.0126)	(0.00431)
Control Mean	0.0753	0.0224	0.0350	0.0204	0.0178	0.136	0.252	0.183	0.0537
Parallel Pre-Trends p-value	0.0105	0.411	0.748	0.903	0.0563	0.0721	0.440	0.869	0.133
Observations	1,399,581	742,375	179,165	179,654	383,556	657,206	146,331	217,010	293,865

Notes: This table reports dynamic estimates of the effect of access to mandated paid sick-leave on the likelihood of an individual being absent from work for care-giving reasons. Estimates are based on the dynamic two-stage differences-in-differences model, where the year before the mandate serves at the reference year. Coefficient estimates for five or more years before and after the mandate are suppressed because they are highly unbalanced. Control means measure average leave-taking among untreated observations and treated observations in the year prior to the mandate. Estimates are reported separately for men (cols 1–4) and women (cols 5–8) based on the age of the youngest child in their household. See also Table 2 Notes.

Table A.5
Effect of Mandated Paid Sick Leave on Own-Illness Work Absences: Role of Children

	Youngest Child Aged $\leq$ 5		Youngest Child Aged 6–18		No Children Aged $\leq$ 18	
	Men (1)	Women (2)	Men (3)	Women (4)	Men (5)	Women (6)
Treatment Effect	0.00442**	0.00398**	0.00794***	0.00685*	-0.00345*	-0.00260
	(0.00201)	(0.00182)	(0.00228)	(0.00392)	(0.00181)	(0.00210)
Control Mean	0.0387	0.0572	0.0386	0.0599	0.0392	0.0558
Parallel Pre-Trends p-value	0.0299	0.000502	0.176	0.0271	0.768	0.106
Observations	179,165	146,331	179,654	217,010	383,556	293,865

Notes: This table reports estimates of the effect of access to mandated paid sick leave on the likelihood of an individual being absent from work for sick leave. Estimates are reported for the full sample, separately for parents whose youngest child is under 6 and whose youngest child is aged 6–18, and for adults with no children under the age of 18 in the house. See also Table 2 Notes.

Table A.6
Effect of Mandated Paid Sick Leave on Own-Illness Work Absences: Dynamic Estimates

		Men				Women			
	Full Sample	Full Sample	Youngest Child Aged $\leq$ 5	Youngest Child Aged 6–18	No Children Aged $\leq$ 18	Full Sample	Youngest Child Aged $\leq$ 5	Youngest Child Aged 6–18	No Children Aged $\leq$ 18
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
4 Years Before Mandate	0.00282**	0.00195	0.00569**	-0.000211	0.00118	0.00385	0.00792	0.00471	0.00135
	(0.00143)	(0.00168)	(0.00240)	(0.00152)	(0.00253)	(0.00244)	(0.00558)	(0.00425)	(0.00264)
3 Years Before Mandate	0.00266***	0.00172**	0.00342*	0.000913	0.00127	0.00375***	0.00604***	0.00203	0.00392**
	(0.000487)	(0.000760)	(0.00186)	(0.00380)	(0.00176)	(0.000943)	(0.00197)	(0.00387)	(0.00167)
2 Years Before Mandate	0.00191**	0.00165	0.000963	0.00369*	0.00105	0.00201	0.000895	0.00182	0.00246
	(0.000905)	(0.00153)	(0.00210)	(0.00216)	(0.00301)	(0.00171)	(0.00210)	(0.00296)	(0.00300)
Year of Mandate	0.00214	0.00471***	0.00712*	0.00299*	0.00439**	-0.00103	-0.0000275	0.00205	-0.00325
	(0.00141)	(0.00148)	(0.00368)	(0.00162)	(0.00218)	(0.00220)	(0.00370)	(0.00344)	(0.00222)
1 Year After Mandate	-0.00130	0.000268	0.00892***	0.00618*	-0.00556**	-0.00324	0.00143	0.000925	-0.00787***
	(0.00188)	(0.00230)	(0.00313)	(0.00326)	(0.00258)	(0.00271)	(0.00442)	(0.00315)	(0.00299)
2 Years After Mandate	0.00391**	0.00241	0.00669*	0.0107***	-0.00248	0.00560**	0.00646	0.0150***	-0.000221
	(0.00152)	(0.00179)	(0.00385)	(0.00302)	(0.00303)	(0.00223)	(0.00398)	(0.00465)	(0.00519)
3 Years After Mandate	0.00389***	0.00483***	0.00219	0.0130***	0.00290	0.00258	-0.00516	0.0145***	-0.000323
	(0.00101)	(0.00163)	(0.00479)	(0.00283)	(0.00213)	(0.00159)	(0.00472)	(0.00326)	(0.00342)
4 Years After Mandate	0.00242	-0.00134	0.00167	0.00774**	-0.00583***	0.00698***	0.0157***	0.00727	0.00254
	(0.00162)	(0.00215)	(0.00834)	(0.00375)	(0.00212)	(0.00220)	(0.00261)	(0.00468)	(0.00202)
Control Mean	0.0433	0.0360	0.0288	0.0427	0.0365	0.0516	0.0470	0.0597	0.0485
Parallel Trends p-value	7.08e-12	0.00212	0.0299	0.176	0.768	0.0000775	0.000502	0.0271	0.106
Observations	1,399,581	742,375	179,165	179,654	383556	657206	146,331	217,010	293,865

Notes: This table reports dynamic estimates of the effect of access to mandated paid sick leave on the likelihood of an individual being absent from work for own-illness reasons for three years before a mandate has passed through three-years after a mandate has passed. See also Table A.4 Notes.

Table A.7
Effect of Mandated Paid Sick Leave on Work Absences: Gender Gaps

	Full Sample	Youngest Child Aged ≤ 5	Youngest Child Aged 6–18	No Children Aged ≤ 18
	(1)	(2)	(3)	(4)
Treatment Effect x Female	-0.00596	-0.00939	-0.00593	0.00241
	(0.00768)	(0.0120)	(0.0126)	(0.00384)
Control Mean	-0.122	-0.222	-0.158	-0.0433
Observations	1,399,581	325,496	396,664	677,421

Notes: This table reports estimates of the difference in the effect effect of access to mandated paid sick-leave on the likelihood of an individual being absent from work for care-giving for women compared to men based on a two-way fixed effects model. See also Table 2 Notes.