The Impact of Job-Protected Leave on Female Leave-Taking and Employment Outcomes

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Abstract

This paper provides evidence on the impact of job-protected family leave on female leave-taking and employment outcomes. I study a state-level paid-leave program in California and exploit the institutional feature whereby, under this program, leave includes job protection only for women who work at firms with 50 or more employees. I find that the increase in leave-taking as a result of California's program is largest for women whose leaves are job-protected. Furthermore, it appears that gains for disadvantaged subgroups (less-educated, unmarried, and minority new mothers) exist only for the subsample of women with access to job protection. I then examine the impact of job-protected leave on female employment. Using a difference-in-difference-in-differences approach, I compare labor-market outcomes for women at large versus small firms in California to women at large versus small firms outside of California after the passage of paid leave. I find that large employers who are forced to offer job-protected leave decrease female hiring by 1.1% in favor of less-costly male employees. However, I also find that female separations decrease by 1.5% as a result of access to job-protected leave, so that female employment overall increases slightly. These results provide evidence of both a supply-side and demand-side effects of job-protected leave. Women are more attached to a labor force that affords them flexibility after childbirth, but are also costlier to employers if they are likely to take leave to care for newborns.

1. Introduction

Two weeks after taking office, President Bill Clinton signed the Family Medical Leave Act (FMLA), which provided eligible mothers 12 weeks of job-protected unpaid leave after giving birth. The passage of this legislation represented a major victory for progressive advocates of family-friendly leave legislation, who had seen two prior iterations of the FMLA vetoed by President George H.W. Bush. Still, the FMLA is viewed by many as incomplete; it is available only to workers at firms with more than 50 employees (meaning over 40 percent of the labor force is uncovered) and does not come with any wage replacement, making it challenging for low-income parents to take advantage of this program (Klerman, Daley, and Pozniak 2012).[1]

The United States is the only developed country in the world that does not offer paid leave to new mothers after the birth of a newborn. In the recent epoch, there has been advocacy on both the state and federal level to provide access to paid parental leave. Many states (California, Hawaii, New Jersey, New York, and Rhode Island) extended the FMLA by providing temporary disability insurance (TDI) for pregnancy-related short-term disabilities. However, pregnancy-related disability leave is not maternity leave to bond with a newborn; in fact, much of it is typically taken before childbirth.

Beyond these TDI programs, a few states have passed legislation providing for access to paid family leave: first, California (passed in 2002 and implemented in 2004); then, New Jersey (passed in 2008 and implemented in 2009); Rhode Island (passed in 2013 and implemented in 2014); New York (to be implemented in 2018); and most recently, DC and Washington (both

[1] Specifically, to be eligible for unpaid FMLA leave, an employee must have worked for her employer for at least one year and accumulated 1,250 work hours. She must work at a location that has 50 employees within a 75-mile radius. "FMLA Eligibility."

to be implemented in 2020) (Rossin-Slater 2017, Corte 2017).[2] State-paid leave in California and New Jersey differs from the federal FMLA in three key ways: (1) unlike FMLA coverage, it is paid, so it provides wage replacement; (2) employees at firms of all sizes are eligible for paid leave; and (3) while FMLA leave is job-protected, paid leave is not—that is, employers are not required to hold jobs open for employees who decide to take paid leave. Yet, when employees take FMLA leave and paid leave simultaneously, their leaves are job-protected. Thus, employees at large firms (50+ employees) in California and New Jersey are eligible for job-protected paid leave, and employees at small firms (fewer than 50 employees) are eligible for paid leave, but without any job protection. In this paper, I compare maternal leave-taking and female employment outcomes for women at firms above the FMLA cutoff to those at firms below the cutoff to isolate the impact of job-protected leave on female labor-market outcomes.

For the purpose of this paper, I focus on the impact of California's Paid Family Leave (CA-PFL) program. I focus on California, rather than paid leave enacted later in New Jersey and Rhode Island for two main reasons. First, the passage and enactment of New Jersey's Family Leave Insurance (NJ-FLI) overlaps with the Great Recession, making the immediate impact of the program challenging to isolate. Second, Rhode Island is excluded because its program is even more nascent than New Jersey's (enacted in 2014) and is structured differently, with job protection provided for female employees regardless of firm size. Further work should compare the impact of Rhode Island's program to the impacts in California and New Jersey. If the

[2] These programs are very similar to TDI in these states, as they are funded by employee contributions. Family leave in these states can start immediately after TDI ends, so we can understand family leave as providing citizens of these states access to longer leaves with higher wage replacement. Another difference between the programs is that TDI applies only to new mothers who may experience childbirth-related disability; whereas paid leave in these states is available to new mothers and new fathers in order to bond with a newborn.

hypotheses in this paper are valid, the difference between small and large firms I document should not exist in Rhode Island.

This paper is novel in its consideration of the interaction between CA-PFL and the FMLA to isolate the impact of access to job protection on female leave-taking and employment outcomes. The contribution of this paper is twofold. I first consider the impact of access to job protection on maternity leave. I demonstrate that CA-PFL doubled the use of maternity leave, increasing it from an average of three to six weeks for new mothers. I then show that these gains are largest for women who work at large firms and thus have access to job protection: average leave-taking after the passage of CA-PFL for the small firm subgroup increases to around 3.3 weeks; average leave-taking for large-firm employees (with access to job protection) is more than twice as long, increasing to nearly 7 weeks. I further establish that gains for disadvantaged (minority, unmarried, and less-educated) new mothers are concentrated in the large-firm subsample. My evidence suggests that access to paid leave is not sufficient for new mothers who cannot afford to take unpaid FMLA leave after the birth of a newborn; instead, job protection is pivotal for these women.[3]

I then consider both supply- and demand-side impacts of CA-PFL on female employment. I use data from the publicly available Census LEHD database, which allows me to explore firm employment across the U.S. and understand industry dynamics and individual employment trajectories. I utilize a standard difference-in-difference-in-differences (DDD) approach, comparing female employment at large versus small firms in California to large versus small firms outside of California to identify the causal effects of the provision of paid leave with job

[3] This finding relates to work by Fass (2009) who highlights the importance of job-protected leave for low-income new mothers and Carneiro et al. (2015), who finds that in Norway, transitioning to paid job-protected leave from a system of unpaid leave has large gains for disadvantaged women.

protection on female labor market outcomes. On the demand side, female hiring at large firms decreased by 0.06 percentage points (1.1%) after the passage of CA-PFL. This is in line with Summers (1989) and Gruber (1992), who caution that group-specific mandates (such as access to job-protected paid leave) that have a disproportionate impact on the behavior of one group (women of child-bearing age) can make this group more expensive to hire relative to their older and male counterparts.[4] But flexible workplace policies have an impact on labor-force attachment as well (Byker 2016). On the supply side, the share of child-bearing-age female separations decreases post-CA-PFL by 0.07 percentage points (1.5%). My separations result is larger in magnitude than the hires result, suggesting the overall impact of job-protected leave on female employment is slightly positive. To evaluate the causal interpretation, I perform a series of robustness checks that suggest my results are in fact attributable to the extension of paid leave in California. When comparing firm sizes unaffected by the FMLA cutoff and evaluating the impact of paid leave on women of non-childbearing age, I find no statistically significant decrease in female hires or separations post-CA-PFL.

The idea that there is both a supply- and demand-side impact of leave legislation is consistent with recent work by Das and Polachek (2015), who note that while labor-force participation increases for young women post-CA-PFL, unemployment duration also increases. My findings are also related to recent work by Bana, Bedard, and Rossin-Slater (2017), who use administrative data from CA-PFL to demonstrate that larger leave benefits increase the

[4] Empirical work by DeLeire (2000) and Acemoglu and Angrist (2001) considers the unintended consequences of group-specific mandates in a different context the Americans with Disabilities Act (ADA) finding a decline in employment for disabled workers after the law's passage that they attribute to increased employer costs of accommodation of disabilities. Jolls (2000) contemplates a puzzle posed by these findings: the costs of providing reasonable accommodation, as mandated by the ADA, are estimated to be small; and yet it has a substantial impact on disabled employment. She suggests that the ADA's other provisions, e.g. its prohibitions on differential treatment of disabled workers, may pose a substantial cost to employers.

likelihood of returning to an employee's pre-leave firm, thus reducing turnover costs for employers.

My conclusion offers some avenues for future research and implications for policy. I demonstrate the importance of access to job protection for family leave to be a realistic option for less-advantaged new parents. Encouraging more men to utilize family leave could make it more difficult for employers to substitute away from leave-likely women. Additionally, educating employers about the benefits of paid-leave extension and encouraging them to develop cost-effective ways to adjust to employee leave-taking will make leaves less costly and thus result in fewer negative externalities.

The remainder of the paper is organized as follows. Section 2 provides additional background on the FMLA and CA-PFL. Section 3 provides detail on the data I use. Section 4 follows the methodology of Rossin-Slater, Ruhm, and Waldfogel (2013) and establishes that increases in leave-taking are largest for women with access to job protection. Section 5 provides suggestive evidence that access to paid leave decreases employers' desire to hire women, while also increasing female labor force attachment. Section 6 concludes.

2. Institutional details

2.1. The U.S. is a leave outlier

The federal FMLA, which requires covered employers to provide employees with job-protected unpaid leave for qualifying medical and family reasons, including the birth or adoption of a newborn, has made it easier for parents to spend time with their children in early life, which is related to superior later-life outcomes.[5] Since its enactment, the FMLA has been used more than 200 million times by women and men who have taken leave to care for their own health or the health of their family members.[6] Still, critics of the FMLA cite various deficiencies with the program. Most glaringly, as discussed previously, the FMLA does not cover employers with fewer than 50 employees and provides access only to unpaid leave, making it difficult for low-income families to utilize the program. It is estimated that nearly two thirds of eligible workers do not take FMLA leave because they are unable to handle the financial burden of unpaid leave (Reagan 2013). Likely stemming from the deficiencies of the FMLA, the extension of paid leave in the United States has become a popular policy issue and one of recent focus for state and local legislators who aim to make the workplace more family-friendly (Miller 2017).

The United States is the only industrialized country in the world without a national system of paid parental leave (Rossin-Slater 2017). A 2014 report by the International Labour Organization of the United Nations surveyed 185 countries and found that only three—the United States, Oman, and Papua New Guinea—failed to offer cash benefits to women during maternity leave.[7] Compared

[5] See Rossin-Slater (2017) for an extensive overview of the impact of family leave policies on children. Some relevant work includes Carneiro et al. (2015) for impact of leave policies on school dropout rates, Ruhm (2000) for impact on child mortality, Stearns (2015) and Rossin (2011) for leave provision reducing low birth weight births, Stearns (2015) finds these concentrated for disadvantaged groups. However some work finds little impact of leave-taking on children in early childhood or on long-term educational attainment or earnings, e.g. Baker and Milligan (2015), Dustmann and Schonberg (2012).

[6] Family Medical Leave Act at 22. (2015).

[7] Maternity and Paternity at Work: Law and Practice Around the World. International Labour Organization. 2014.

to the U.S., the rest of the developed world's paid leave offerings are extremely generous. Canada provides for one year of paid leave, with around 55% of wages replaced, and European countries provide up to 15 weeks of maternity leave with between 70% and 100% of wages replaced (Addati, Cassirer, and Gilchrist 2014). In the Middle East, leaves are typically shorter (for a period of on average 14 weeks), but come with 100% wage replacement for women during this period.[8] Paternity leave is more limited; still, about 50% of countries offer access to paid leave for new fathers, and this leave is almost always paid (Rossin-Slater, Ruhm, and Waldfogel 2017).

Given the interest in flexible workplace policies and the fact that the U.S. is an outlier in the parental-leave space, proponents of a federal program of paid leave have emerged in the private sector and political and academic spheres. Industry leaders like YouTube CEO Susan Wojcicki[9] and Facebook founder Mark Zuckerberg, among many others, have campaigned for access to paid leave and for paternity leave benefits.[10] In the 2016 presidential election cycle, Democratic candidate Hillary Clinton put forth a plan to guarantee up to 12 weeks of paid leave and ensure a two-thirds wage-replacement rate for workers,[11]and then-candidate Trump promised access to six weeks of paid maternity leave for new mothers.[12]

2.2. State-level paid leave in California

CA-PFL offers six weeks of partially paid leave to bond with a newborn or a recently placed foster or adoptive child, or to care for a sick family member (Rossin-Slater, Ruhm, and Waldfogel 2013). For bonding claims filed to spend time with a newborn, the six weeks of leave must be

[8] An outlier here is Syria, which meets ILO standards by providing 120 days of leave to new mothers and 100% of earnings during this period, though this generous leave is offered only for the birth of the first child. Maternity and Paternity at Work: Law and Practice Around the World. International Labour Organization. 2014.

[9] Wojcicki (2014).

[10] Zuckerberg took two months of paternity leave to spend time with his own newborn, see Bravo (2015).

[11] Clinton (2015).

[12] But no similar provision for new fathers, Sullivan and Costa (2016).

taken within the first year of the child's birth, adoption, or foster placement.[13] CA-PFL operates through the state's short-term disability insurance (SDI) program; thus, funding comes entirely from employee contributions.[14] Unlike the relatively limited coverage under the FMLA, almost all private-sector workers are eligible for CA-PFL. Wage replacement with CA-PFL is far from complete, but it is substantial, providing 55% of a worker's wages (up to a ceiling based on the state's average salary) (Rossin-Slater, Ruhm, and Waldfogel 2013).

Women are much more likely to take parental leave, and take much longer leaves, than men; as such, the impact of leave programs on labor-market outcomes is more pronounced for women of childbearing age than for their male (or older female) counterparts (Bartel et al. 2015). In 2013, 14% of employed mothers with a child under the age of one reported being on leave, compared to only two percent of employed fathers with a child under the age of one (Bartel et al. 2015).

Despite the fact that employees pay through deductions for paid-leave provisions, awareness (and consequently use of paid leave offerings) remains surprisingly low. A 2009–2010 survey found that fewer than half of Californians who recently experienced a qualifying event under CA-PFL knew of the program's existence, and awareness is lowest for workers who earn less than $30,000 annually (Engeman 2012). Even among those with awareness of paid family leave, usage may be limited because of the lack of job protection for workers under CA-PFL. However, employees who are covered by the FMLA (at firms with more than 50 employees in a 75-mile radius of their worksite) must take CA-PFL and FMLA leave concurrently, and thus receive job protection even for their paid leaves.

The provision of job-protected leave increases the firing cost of employers' at large firms by precluding the dismissal of employees for taking costly leaves. This distinction is key to my

[13] "Paid Family Leave Factsheet." Employment Development Department. State of California.
[14] Ibid, the same is true for New Jersey and Rhode Island's programs.

identification strategy, as I test the impact of CA-PFL on leave-taking and employment outcomes for new mothers at firms above and below this 50-employee FMLA cutoff. Understanding precisely what is meant by job protection and how its unavailability may impact employee and employer behavior is thus key to my analysis.

In the context of the FMLA, if employees believe they were terminated for taking leave, they are permitted to sue their employers through an FMLA inference claim. While the specifics of the case law vary in different circuits, broadly, for an FMLA inference claim to stand, employees must prove that 1) they were eligible for FMLA leave, 2) an adverse action by their employer limited their ability to take FMLA leave, and 3) the adverse action was precipitated by their taking of leave or request to take leave (Brown v. ScriptPro, LLC. 700 F. 3d 1222). An employer's intent is not directly relevant to the interference analysis (Grace v. USCAR, 521 F.3d 655, 670. 6th Cir. 2008), but if an employer can proffer a legitimate business reason for an employee's termination that is unrelated to interfering with FMLA rights, the burden shifts to the employee to refute this reason (Donald v. Sybra, Inc. 667 F.3d 7 757 762. 6th Cir. 2012). As such, it is an overstatement to suggest that workers cannot be fired while taking FMLA leave. As long as the termination is for a legitimate business reason unrelated to leave-taking, it is lawful. Practically, though, in firing workers while on leave (or who have given notice that they will take FMLA leave), employers are exposing themselves to costly litigation.

The same is not true for paid leave. Workers can be dismissed for taking (or announcing an intent to take) paid leave when their jobs are not federally protected by the FMLA. Thus, around 40% of California's workers could be fired for taking CA-PFL. Unsurprisingly, the threat of job loss influences employee leave-taking patterns. A 2011 survey found that 37% of workers who knew about the existence of PFL decided not to take it because they were worried that their

employer would be unhappy, that their opportunities for advancement would be affected, or that they might actually be fired.[15]

This distinction is key to my study. First, it is clear that the provision of job protection will make paid leave more expensive for larger firms relative to their smaller counterparts because (1) more employees will take leave (and often for longer periods) without the worry that they can be lawfully terminated for their decision to do so;[16] and (2) thus, employers will have to hold jobs open for workers who take leave—for example, by hiring and training a temporary worker or paying current employees a premium for working extra hours. The employer cost of CA-PFL is illustrated by opposition to a recent bill in the California state legislature that would have extended job protection under CA-PFL to businesses with more than 20 employees. California Governor Jerry Brown ultimately vetoed the bill in the face of tremendous opposition from small businesses: the California Chamber of Commerce labeled the proposed bill a job killer for the reasons discussed above (Barrera 2016).

3. Data and Methodology

3.1. Impact of CA-PFL on female leave-taking

3.1.1. Data: March Current Population Survey

I first study the impact of CA-PFL on maternal leave-taking, following Rossin-Slater, Ruhm, and Waldfogel (2013), who use data from the March Current Population Survey (CPS) Annual Demographic Supplement, accessed via the Integrated Public Use Microdata Series (IPUMS) database. The March CPS provides detailed information on leave-taking and labor market outcomes for a large and nationally representative sample. Since 1994, survey respondents have

[15] Appelbaum and Milkman (2011).

[16] This is consistent with the leave results documented by Bartel et al. (2015) who estimate that CA-PFL increases fathers' leave-taking by approximately 2.4 days and mothers' leave-taking by 6 days.

been able to indicate that they were absent from work in the prior week because of maternity/paternity absence. An important limitation of the March CPS is it does not provide precise information on child-birth dates and women's employment status during pregnancy. I am able to observe only whether a new parent in the labor force was absent from the labor force in the prior week because of family leave; not whether new mothers drop out of the labor force in advance of giving birth. Importantly for my analysis, since 1988 the March CPS supplement has included a question on firm size. This allows me to ascertain whether increases in leave-taking are driven by the subsample of workers who have access to paid leave with job protection (employees at firms with 50+ employees) relative to those who are below the FMLA cutoff and thus do not have access to job-protected paid leave.[17]

Like Rossin-Slater, Ruhm, and Waldfogel (2013) I analyze leave-taking using a difference-in-difference approach and a variety of dependent variables. The baseline leave variable classifies women as leave-takers if they report being on maternity leave in the prior week. However, because new mothers (and fathers, although not the focus of my analysis) may be on CA-PFL leave and misreport on the March CPS, I also construct a broad family-leave variable that characterizes a new parent reporting time off work for vacation/personal days and child-care and other family-related reasons as a leave-taker. As a placebo, I provide results for only "other" leave, which includes all of the broader leave responses except for maternity/paternity leave, to determine whether any of the impact of CA-PFL I document is driven by in-leave-taking generally rather

[17] Like Rossin-Slater, Ruhm, and Waldfogel (2013) I use the sample period of 1999-2010 in e orts to make my results as comparable to theirs. Importantly, the firm size question on the March CPS is known to be a problematic one, as the Department of Labor has acknowledged. Many respondents provide answers about the establishment or location of work rather than the overall size of the rm. Additionally, prior to March 2011 there were firm size partitions at 10 or fewer, 25 or fewer, and 100 or fewer employees, rather than 10, 50, and 100, as is the new standard. As such, I compare leave-taking by survey respondents who work at small firms (25 or fewer employees) to large firms (100+ employees). The lack of precision in the data biases against my finding differences in the small and large firm subsample because employees who report working at a small firm may in fact work at small establishments that are part of large firms required to provide access to job-protected leave.

than maternity leave specifically. The main specification involves comparing leave-taking by women in California whose youngest child was born within the last year to leave-taking by women in California whose youngest child is aged 5–17. Unfortunately, the March CPS has not consistently asked about firm-size categories directly above/below the 50-employee FMLA cutoff. Thus, I label women at firms with fewer than 25 employees as small-firm employees and those with more than 100 employees as large-firm employees.

3.1.2. Methodology: DD approach comparing leave-taking by new mothers in CA to various control groups[18]

I follow Rossin-Slater, Ruhm, and Waldfogel (2013) and employ a difference-in-difference design, comparing changes in the outcomes for California women with infants surveyed before and after the implementation of California's paid family leave. My baseline empirical strategy involves estimating the regression specification below for (1) the full sample, replicating the results of Rossin-Slater, Ruhm, and Waldfogel (2013); (2) the large-firm subsample; and (3) the small-firm subsample. The baseline regression estimated is

$$Y_{i,t} = \beta_0 + \beta_1 * Treat_i + \beta_2 * Post_t * Treat_i + \gamma' X_{i,t} + \delta_t + \epsilon_{i,t}$$

for an individual surveyed in year *t.* $Y_{i,t}$ is the dependent variable of interest, equal to 1 for the use of leave (for various definitions); $Treat_i$ is a dummy variable equal to 1 for California mothers whose youngest child is under one year of age on the survey date; $Post_t$ is a dummy variable equal to 1 if the individual was surveyed in the year following the implementation of CA-PFL (2005 or later) and 0 otherwise; $X_{i,t}$ is a set of individual-level controls for standard demographic

[18] The first part of my paper extends Rossin-Slater, Ruhm, and Waldfogel (2013), extending the author's work by estimating their specifications separately for the large and small firm subgroup. As such, the methodology I adopt below is exactly the methodology they implement in their work, and I adopt their notation for ease of comparison.

characteristics: marital status, race, education, age, and birthplace; and δ_t is a vector of general year-effects. The coefficient of interest β_2 measures the DD estimate of CA-PFL on the treatment group of new mothers in California.

The key identification assumption in this DD strategy is that changes in the leave outcomes would have been the same for treatment and control groups if CA-PFL had not been implemented. This is a fundamentally untestable assumption; to address it, I include a variety of control groups: (1) women in California with a youngest child between the ages of 5–17 (the main specification of interest); (2) women in California with no children; (3) women with young children in the next three-largest states: Florida, New York, and Texas; and (4) mothers with infants residing in all states other than California. As a placebo and further check the veracity of this assumption, I obtain a set of DD estimates for other states with temporary disability insurance programs (Hawaii, New York, and Rhode Island) that did not implement paid family leave policies during the treatment period examined (1999–2010). These states offer some access to paid leave, and if I am simply capturing differences in leave-taking that are correlated with the period when CA-PFL was passed, I would expect to see increases for California, as well as the other TDI states.

The estimates of the impact of CA-PFL are potentially biased if patterns of fertility (that is, typical ages of new mothers) are impacted by CA-PFL or if CA-PFL induces migration into California (or into large firms relative to small firms, which are forced to offer job-protected leave through the FMLA) to take advantage of the existence of (job-protected) paid leave. I believe it is unlikely that the behavioral impact of CA-PFL is a significant contributor to the increased leave-taking documented, especially given that awareness of the benefit program, even for the subsample of employees eligible in California, is fairly low (Applebaum and Milkman 2011). In Table 1, I present summary statistics for employment composition overall, and at large and small firms

separately, in California for my treatment group of mothers who gave birth in the last year.[19] There is little evidence of any drastic changes in demographic characteristics of employees as a result of the passage of CA-PFL.

For these DD estimates, it is not possible to cluster standard errors at the treatment group-level, because most specifications contain only a control and treatment group within the same state (California). As an alternative, I follow the two-step method developed by Donald and Lang (2007) and utilized by Rossin-Slater, Ruhm, and Waldfogel (2013) to account for potential serial correlation in the error structure. As a first step, I calculate regression-adjusted differences in outcomes between the treatment and comparison groups in each survey-year:

$$Y_{i,t} = \gamma' X_{i,t} + \pi_t' * Treat_i + \gamma_t + \epsilon_{i,t}$$

for each individual *i* surveyed in year *t*. The vector π_t' contains regression-adjusted differences between the treatment and comparison groups for each survey-year. In the second stage, I collapse the differences into survey-year cells and estimate

$$\pi = \rho_0 + \rho_1 * Post_t + u_t$$

where the key coefficient of interest ρ_1, if non-zero, implies that the differences in leave-taking between the treatment and control group changes after the implementation of CA-PFL. Because the second stage of Donald and Lang in this case is a regression with only 12 observations, I use a t-distribution with 10 degrees of freedom instead of a standard normal distribution to test for statistical significance.[20]

For models with the control group as mothers in all other states, I am able to compare the coefficients and standard errors obtained using this DL method to those from the DD method with

[19] Table A1 has much more extensive descriptive statistics and also includes summary information for the other control groups studied.

[20] I use March CPS Supplement person weights for each survey year.

robust standard errors clustered at the state level, and obtain comparable results with both approaches. I report the latter.

3.2. Impact of CA-PFL on Female Labor Market Outcomes

3.2.1. Data: Census Longitudinal Employer-Household Dynamics Quarterly Workforce Indicators

I then use data from the publicly available Census Longitudinal Employer-Household Dynamics (LEHD) partnership to estimate the causal impact of CA-PFL on female employment. LEHD integrates data from state-supplied administrative records on workers and employers with censuses, surveys, and other administrative databases to provide longitudinal data on US employment patterns.[21] The primary data for my analysis are from LEHD's flagship product, the Quarterly Workforce Indicators (QWI), which provides information about trends in employment, hiring, job creation, and destruction, with data on both firm and employee characteristics, that is, firm size and employee race, gender, and educational background. This data is at a national level. While unfortunate, I do not think it argues against my findings because (1) it biases them downward, since I am comparing firms that are definitely below the FMLA cutoff to a mix of firms above and below and (2) more than 95% of firms have only one establishment, and those that do not are super-large firms that I do not consider in this paper (BLS). [22,23] The LEHD data on firm size are preferable to the March CPS because the firm size buckets allow for focus on firms just below (20–49 employees) and firms directly above (50–249 employees) the FMLA cutoff. However, these data are still limited by the fact that I am not able to observe firm size directly. An

[21] Local Employment Dynamics (2015)

[22] This data is at a national level. While unfortunate, I do not think it argues against my findings because (1) it biases them downward, since I am comparing firms that are definitely below the FMLA cutoff to a mix of firms above and below and (2) more than 95% of firms have only one establishment, and those that do not are super-large firms that I do not consider in this paper (BLS).

[23] Local Employment Dynamics (2015).

ideal test would allow the use of firm-level micro-data to construct more helpful groupings; unfortunately, the publicly available data do not allow for this level of granularity.

Because data are reported at the firm size–quarter level, the QWI allows me to determine whether the impact of CA-PFL is different in firms above and below the 50-person cutoff. Additionally, QWI publicly available data allow for industry-level analysis. This is relevant because recent work (for example, Bartel et al. 2015) has suggested that there are differential patterns of leave-taking for different industries/occupations, finding, for example, that fathers in occupations with a high share of female workers are more leave-likely. QWI industry-level data also facilitate the inclusion of industry fixed effects.

The publicly available LEHD data have important limitations. While allowing for investigation of firm employment patterns based on employee (1) sex and education, (2) sex and age, and (3) sex and race, I am not able to examine these employee characteristics simultaneously. Based on both theory (Gruber 1992, 1994) and past empirical work on this topic (Ross 2013), I hypothesize that on the demand side, paid leave most negatively impacts women of child-bearing age who have low skills and/or are less educated, as they are most substitutable. However, I am not able to disentangle the impact on high-skill women of child-bearing relative to their low-skill counterparts. Additionally, the publicly available LEHD data become available for states in staggered intervals. I focus on the sample period of 2000–2010, with 2000–2004 as the pre-CA-PFL period and 2005–2010 as the post-period. Because my pre-PFL period begins in 2000, I exclude those states that do not have data available until after 2000. Excluded states are Alabama, Arkansas, Arizona, DC, Kentucky, Massachusetts, Mississippi, New Hampshire, and Wyoming. I also exclude New Jersey because of its implementation of paid leave toward the end of my sample period.

3.2.2. Methodology: DDD approach comparing labor market outcomes for small versus large firms in California to small versus large firms outside of California

To identify the impact of CA-PFL on female labor-market outcomes, I compare small and large firms in California that were impacted by the passage of CA-PFL to small and large firms in other states that were not impacted. The DDD specification will allow for a causal interpretation as long as changes in female employment outcomes at large versus small firms in California and other states would have been the same but for the passage and enactment of CA-PFL. Of course, this assumption is untestable, but I provide some support for it in Figure 1, which documents the following pre-trends for each of my dependent variables: (1) the fraction of female employees at large versus small firms; (2) the fraction of female hires at large versus small firms; and (3) the fraction of female separations at large versus small firms.

To make the results comparable to allow for an estimate of overall employment impact, I compute female hires and female separations as a share of total employees. In any particular quarter, I have

$$E_{m,t} + E_{f,t} = E_{m,t-1} + E_{f,t-1} + H_{m,t} + H_{f,t} - S_{m,t} - S_{f,t}$$

and my dependent variables ($E_{i,s,t}$ below) are $\frac{E_{f,t}}{E_{m,t}+E_{f,t}}; \frac{H_{f,t}}{E_{m,t}+E_{f,t}}; \frac{S_{f,t}}{E_{m,t}+E_{f,t}}$ where $E_{f,t}$ is female employment in period t, $H_{f,t}$ is female hires in period t, and $S_{f,t}$ is female separations in period t.

Following the methodology of Garthwaite, Gross, and Otidid, (2014), my triple-difference analysis compares employment decisions by large versus small firms within California to large versus small firms outside of California, before and after the passage of CA-PFL. This triple-difference regression specification is

$$E_{i,s,t,g,z} = \gamma_i \cdot \delta_s + \gamma_i \cdot \varphi_t + \gamma_i \cdot \pi_g + \gamma_i \cdot \theta_z + \delta_s \cdot \varphi_t + \delta_s \cdot \pi_g + \delta_s \cdot \theta_z + \pi_g \cdot \varphi_t + \pi_g \cdot \theta_z + \varphi_t \cdot \theta_z +$$
$$\beta \cdot I\{z = Large\} \cdot I\{s = isCA\} \cdot \{t = Post(\geq 2002)\} + \epsilon_{i,s,t,g,z}$$

where $E_{i,s.t,g,z}$ is defined above (I also estimate the above for $H_{i,s,t,g,z}$ and $S_{i,s,t,g,z}$), $Post_t$ is an indicator that takes value 1 after the passage of paid leave in California in Q3 2002,[24] $Large_z$ is an indicator equal to 1 for a firm with between 50–249 employees and 0 for a firm with 24–49 employees, and $isCA_s$ is an indicator for the observation from California. I include industry (γ_i), geography (δ_s), year-quarter (φ_t), age group (π_g), firm size (θ_z) fixed effects, and all of the two-way interactions between these sets of fixed effects. The main coefficient of interest is that on the triple interaction term, β. This coefficient represents the effect of paid leave on female labor market outcomes for large firms in California that were required to offer job-protected paid leave after the passage of CA-PFL. The presence of my full set of fixed effects makes it unnecessary to include the main effects of *Large*, *isCA*, and *Post* (and their two-way interactions) explicitly in the regression.

4. Results on leave-taking following Rossin-Slater, Ruhm, and Waldfogel (2013)

4.1. Leave-taking increases most for women at large firms with access to job protection

Table 1 provides summary statistics for the treatment group (California mothers with youngest child under one year in age) and the main control group (California mothers of youngest children aged 5–17) before (1999–2004) and after (2005–2010) the passage of CA-PFL.[25] All statistics are weighted by the CPS March Supplement person-weights. The pre-PFL treatment group mean for

[24] Results reported below are qualitatively similar for enactment (Q2 2004) as marking the post period. I use the passage of CA-PFL as the main date of interest to allow for employee/employer adjustment even before the law is implemented. In the leave-taking results above using the date of passage would not be sensible because women were not entitled to protected leave until the legislation was enacted. As such, it is hard to imagine why leave-taking would increase in the period following the passage of CA-PFL but before its implementation.

[25] Appendix Table A1 replicates the full set of Rossin-Slater, Ruhm, and Waldfogel (2013) summary statistics. While not identical to their estimates, my summary statistics for the whole sample are very close to theirs. For brevity, here I report only a few leave categories as well as summary statistics for the disadvantaged subgroups I focus on.

previous-year employment is 59.1%, which is the estimate of the share of female employees who would have been eligible for CA-PFL leave.

Notably, there is a significant jump in maternity leave for the treatment group post-2004 relative to the pre-period (from 5.3% to 11.4%)[26] Not surprisingly, there is essentially no maternity-leave use by the control group of women whose youngest child is between the ages of 5–17 in either period. In Panel B and Panel C, I consider large- and small-firm employees, respectively. Comparing these two panels provides the first evidence that leave-taking patterns are influenced by the provision, or lack, of job protection. Only 2.4% of women at firms with fewer than 25 employees reported being on maternity leave in the pre-period (an average of 52*2.4% weeks of leave for each new mother, or 1.2 weeks), compared with over 7% of eligible women at large firms (an average leave of more than 3.6 weeks). Both groups report increases in leave-taking as a result of the passage of CA-PFL. For mothers of infants at small firms, who are entitled to unprotected paid leave in the post-period, leave-taking rises to about 7% in the post-CA-PFL period (approximately the pre-PFL mean for large-firm employees). For mothers of infants at large firms, leave-taking jumps from 7.1% in the pre-period to 13.5% (to more than seven weeks of leave on average) after the passage of CA-PFL.

Table 2 presents regression results for the three leave outcomes, using as a control California mothers whose youngest child is between the ages of 5–17. Panel B and Panel C report results for the large- and small-firm subgroups, respectively.[27]

[26] This is very similar to the Rossin-Slater, Ruhm, and Waldfogel (2013) estimate, which rises from 5.3% to 11.8% in the post period.

[27] In Appendix Table A2 I replicate Rossin-Slater, Ruhm, and Waldfogel (2013)'s baseline result, including ITT specifications that do not condition on prior year employment. Such an exercise is not sensible for the large and small firm subsamples, as an individual reporting a firm size was necessarily employed in the prior year. In Appendix Table A3, I present results for the whole firm sample and for the large and small firm subgroups for each of the control groups Rossin-Slater, Ruhm, and Waldfogel (2013) consider: (1) mothers of youngest children aged 5 to 17 in California; (2) women with no children in California; (3) mothers of youngest children aged less

Like Rossin-Slater, Ruhm, and Waldfogel (2013), I find a large and statistically significant increase in leave-taking for women as a result of the passage of CA-PFL. Access to paid leave through CA-PFL increases female leave-taking by a statistically significant 5–6.5 percentage points, more than doubling the use of maternity leave—from an average of about three weeks to six or seven weeks for new mothers. Unsurprisingly, the magnitudes of the PFL effects on family leave, and any leave, are similar to those estimated for the maternity leave. As a placebo test, I consider the impact of the passage of CA-PFL on other types of leave (excluding those who respond that they are on maternity leave in Column 3) and see no increase as a result of CA-PFL, which bolsters the causal interpretation by suggesting that the effects captured are not attributable to changes in the overall leave landscape.

Panel B reveals larger gains for women in the large-firm subsample than for women overall. Leave-taking for large-firm employees increases by 6.4 percentage points (3.3 weeks), to nearly seven weeks in total, and for small-firm employees by 4.6 percentage points (2.4 weeks), making average leave post-CA-PFL about half as long for small-firm new mothers compared to their large-firm counterparts. Stated another way, the extension of paid leave does not bridge the pre-CA-PFL gap in leave-taking for large- and small-firm new mothers. Many women at small firms still cannot afford to take advantage of CA-PFL, even though this leave is paid, because of fear of losing their jobs for taking unprotected leave.

In my baseline specification, I find that the passage of CA-PFL increases the likelihood of maternity leave for a new mother from around 7.1% in the pre-period to around 13.5% in the post-

than 1 in Florida, New York, and Texas; (4) mothers of youngest child aged less than 1 in all states except California. Interestingly, the small firm increases in leave-taking disappear for some of the broader control groups considered. This suggests that the increase in leave-taking documented for small firm employees as a result of the passage of CA-PFL may be attributable to time-varying trends in leave-taking for small firm employees across the country that coincided with the passage of CA-PFL.

period. This raises leave-taking from an average of around 3.7 weeks (7.1% x 52) in the pre-period to around 7.02 weeks in the post-PFL period. Unsurprisingly in Panel C, women at small firms (fewer than 25 employees) who are not eligible for job protection under the FMLA are much less likely to take maternity leave in the pre-period (around 2.4% probability of leave, which translates to an average leave of about 1.2 weeks in the pre-CA-PFL period). Post passage of CA-PFL, maternity leave rises for this group by around 4.6 percentage points. However, in terms of total magnitudes, even after the passage of CA-PFL, women at small firms take only about 3.5 weeks of leave. Two points are worth noting. First, the FMLA's guarantee of unpaid leave made it much more likely that eligible women would be able to take maternity leave to bond with a newborn. The baseline differences in leave-taking for large versus small firms in California per CA-PFL illustrate this. Second, the extension of paid leave to all California employees did not bridge this gap. Women at large firms still take about twice as much leave as small-firm employees. I attribute this continued difference to the access to job protection afforded to large-firm employees, but not to small-firm employees.[28]

[28] In Table A4, I replicate the Rossin-Slater, Ruhm, and Waldfogel (2013) placebo tests, looking to see whether the measured impact on leave-taking of CA-PFL could be attributed to overall changes in the leave-taking landscape. To test for this, I label as the treatment group three states (Hawaii, New York and Rhode Island) that did not have a paid leave program during this time period but did provide some access to paid leave through their Temporary Disability Insurance programs. The near-zero and statistical insignificance of these coefficients suggests that changes in leave-taking more broadly are not responsible for the impact of CA-PFL that is documented. What is most interesting about these falsification tests is what happens to the magnitude of these coefficients when I break out the small and large firm subsamples separately. For small firms (Panel C) for the baseline control group (women outside TDI states with a child aged <1). I see a statistically significant increase in leave in the placebo TDI states that is of comparable magnitude to the estimated increase in leave-taking as a result of CA-PFL documented in Table 2. And for all the comparison groups, the coefficients remain large for the small-firm subsample, but are much smaller for the all small- and large-firm samples, perhaps suggesting that my estimates of the increase in leave-taking for small-firm employees as a result of CA-PFL may be inflated.

4.2. CA-PFL increases leave-taking for less advantaged subgroups only when leaves are protected

As other authors have noted, when paid leave replaces a system of unpaid leave, increases in leave-taking are largest among disadvantaged mothers (unmarried, less-educated, and minority mothers) who were least able to take advantage of unpaid leave programs (Carneiro, Loken, and Salvanes 2015; Appelbaum and Milkman 2011), because women who cannot afford unpaid leave are unable to take it. It is sensible to consider whether, even when a program of paid leave is implemented, the lack of job protection makes it difficult for some less-advantaged subgroups to take advantage of this leave offering.

Differences exist in baseline leave-taking (prior to CA-PFL) among different educational and ethnic groups. These differences are reported in Table 3, Panel A. In the pre-period, only 2.2% of women in the treatment group with a high-school education or less reported being on maternity leave, compared to more than 9% of women with college or post-college degrees. And only 2.1% of black mothers reported being on leave in the pre-period, compared to 6.9% of white mothers. Clearly, the CA-PFL increases leave-taking for all of the disadvantaged subgroups. Gains are largest for black women, for whom leave-taking rises by an estimated 9.9 percentage points (a 5.15-week increase in leave as a result of CA-PFL). Leave-taking for Hispanic mothers increases as well, by around 5.06 percentage points, or a just over 2.5 weeks. Interestingly, there are no statistically significant increases in leave for women with a college degree or higher, suggesting that the impediment for leave for this group was not that it was unpaid in the pre-PFL period.

In Panel B and Panel C of Table 3, I consider the impact on leave-taking for disadvantaged groups in large- and small-firm subsamples separately. I first note that the differences in leave-taking for large- and small-firm employees are pronounced for most subgroups. Of women with

some college education in the large-firm subsample, 7.8% are on leave in the pre-CA-PFL period, compared with only 1.5% of women with some college education in the small-firm subsample. Similarly, 5.4% of Hispanic new mothers at large firms are on leave in the pre-CA-PFL period, compared to only 1.9% of Hispanic new mothers at small firms.

I find that all of the gains in leave-taking as a result of CA-PFL for disadvantaged new mothers are concentrated in the large-firm subsample (Panel B). The only subgroup for which leave-taking increases in the small-firm subsample is married mothers. In the case of uneducated, unmarried, and minority mothers, the DDD point estimates for the small-firm subsample (Panel C) are small and even negative (for black new mothers), and they are never statistically significant. The results for the large-firm subsample are economically and statistically significant, indicating increases in leave-taking of four to six weeks for women with a high school education or less, women with some college, unmarried mothers, and Hispanic mothers. Interestingly, there are no gains for large-firm employees with a college degree or higher (as for the whole sample, suggesting lack of pay is not the impediment to leave for this group) or white mothers. Likely due to the small sample size, the results for black mothers (for whom overall gains are largest in Panel A) are no longer statistically significant for the large-firm subsample, although the coefficient on the DDD remains large. These results suggest that job protection is key to make leave-taking possible for less-advantaged mothers.

5. Impact of CA-PFL on female labor market outcomes

Recognizing that the increase in leave-taking as a result of CA-PFL is concentrated in large firm-employees, I next test whether CA-PFL has a differential impact on female employment at large versus small firms. Using the logic of Summers (1989) and Gruber (1992), since it is perhaps difficult for wages to adjust for the fact that men are less leave-likely than women (for example,

because of anti-discrimination laws), I would expect female hiring to decrease at large firms relative to small firms. Furthermore, if women are more attached to a labor force that is more exible (Bana, Bedard, and Rossin-Slater 2017), on the supply side, it is likely that female separations should decrease as a result of the passage of paid leave. The sign of the overall employment effect will depend on the relative magnitude of impact of CA-PFL o female hires and separations.

Table 4 reports summary statistics for the sample of firms used in my preferred empirical specification, which compares employment patterns in California to firms in non-treatment states. Annual averages are also provided for the fraction of firm employees that are female, the fraction of hires that are female, and the fraction of separations that are female (the three dependent variables of interest) for all firms, and then large (50–249 employees) and small (20–49 employees) firms separately. In the Appendix, Figures A1 and A2 provide additional descriptive information, presenting data on the age distribution of employees in the pre-CA-PFL and post-CA-PFL period; and the concentration of female employees across the country in these periods.

5.1. Job-protected leave decreases share of female hires and separations

Table 5 shows the estimated effect of paid leave on female employment outcomes in California in the preferred DDD specification. Here, I compare employment outcomes for women in small and large firms in California (who were impacted by paid leave) to small and large firms outside of California.[29] As I note above, to compute my dependent variables, I divide female employees, female hires, and female separations each by total employees in each quarter, thereby making the three coefficients I compute in my forthcoming analysis comparable.

[29] I exclude from the sample New Jersey, because of its consideration of and eventual passage of paid leave toward the end of this sample.

The passage of paid leave with job protection led to a 0.06 percentage point decrease in the share of female hires at large versus small firms in California. Relative to the sample mean of 5.2% of current quarter employees being new female hires, this decrease corresponds to a 1.1% decrease in female hiring at large firms, where paid leave is job protected, relative to small firms, where leave is unprotected. I also see a statistically significant decrease in the share of this quarter's employees who separate from large relative to small firms after the passage of paid leave. The passage of CA-PFL led to a 0.07 percentage point decrease in the share of female separations at large versus small firms, or a 1.5% decrease in female separations. Overall, it appears that the decrease in female separations is of a slightly larger magnitude than the decrease in female hiring, and in aggregate, the passage of CA-PFL leads to a statistically significant 0.65 percentage point (1.4%) increase in female employment.

5.2. Robustness checks

To interpret the DDD estimate as the causal effect of the provision of paid leave on female employment, the implementation of paid leave cannot be correlated with time-varying determinants of female labor-market outcomes in large versus small firms in the sample. That is, to estimate the impact of CA-PFL with my DDD approach, I assume that, absent CA-PFL, there would have been no difference in how the share of female employees, hires, and separations would have evolved at large versus small firms inside of California relative to large versus small firms outside of California. This is an untestable assumption. I document the pretrends for my DDD specification in Figure 1. While the pretrends are imperfect, there does appear to be evidence of a decrease in all three of my dependent variables of interest around the same time as the passage of CA-PFL. These pretrends also suggest that employers and employees began to adjust to CA-PFL

even before its enactment: The downward trajectory for the dependent variables begins around the time of the proposal of CA-PFL in late 1999.

I may also be concerned that the passage of CA-PFL induced selection into the treatment group because leave-likely employees opted into large firms once they realized CA-PFL provided better benefits in large firms compared to their smaller counterparts. It is also a concern that firms, aware of the differential requirements around the 50-employee cutoff, sought to remain just below the cutoff so as not to adhere to the job protection requirements. This concern would be best alleviated by gaining access to LEHD micro data on firm size to test for bunching just below the FMLA cutoff. In the absence of microdata, placebo tests provide evidence for the baseline causal interpretation.

5.2.1. Placebo tests for different firm sizes

One concern is that the baseline results are capturing differential trends in female hiring based on firm size that are fully independent of the impact of CA-PFL. It is possible that larger firms in California have become less likely to hire women, and that current female employees in California at large firms have become less likely to separate from them for reasons unrelated to CA-PFL, and this could affect my results. I address this hypothesis with a series of placebo tests in Table 6. Here I replicate the baseline DDD analysis from Table 5, now comparing small (20–49 employees) and smaller firms (0–19 employees), neither of which must offer job protection post-CA-PFL, and large (50–249 employees) and larger (250–499 employees), both of which must offer job protection. If the baseline results are driven by some California-specific trends, I would expect to capture these same results in the placebo regressions. Table 6 demonstrates that this is not the case.

In my baseline results, the share of this quarter's employees that are female hires and female separations decreases at large relative to small firms as a result of the passage of paid leave. In the

placebo tests in Table 6, the opposite is true. The share of employees that are female hires or female separations increases at large relative to smaller firms in both placebo groups. This suggests that the hiring and separation results we observe in California as a result of CA-PFL are hard to attribute to general trends in female hiring and separations at large relative to small firms in this period.

The placebo employment results are slightly perplexing (female employment increases for workers at firms with 20–49 employees relative to firms with 0–19 employees, and decreases for firms with 250–499 employees relative to firms with 50–249 employees). However, these results are barely statistically significant (significant at the 10% level, relative to the much stronger results for hiring and separations). Furthermore, they do not suggest that systematic increases in female employment at large versus small firms are driving my results, as the coefficients change sign depending on the placebo treatment group considered.

5.2.2. Placebo tests for older workers

A related concern is that the results attributed to CA-PFL are in fact related to changes in labor-market outcomes for women generally and unrelated to the enactment of CA-PFL. To address this concern, I use a placebo age-group test, replicating the baseline DDD results for the older sample of female employees (those aged 45+). If my results are driven by the fact that the passage of CA-PFL coincided with a general tendency for large firms to hire fewer women relative to small firms in California (or relatedly, for women to separate less often from large versus small firms around this time), I should still see a statistically significant decrease in female hiring and separations for this older employee sample. If instead, CA-PFL is responsible for the hiring and separation results documented above, I would not see a statistically significant difference in female hiring, separations, or employment for older women who are unlikely to take maternity leave. In Table 7, I find no statistically significant impact on female employment or separations for the older

subsample. I do see a statistically significant increase in female hiring of 0.031 percentage points, or 1.8% relative to the pre-PFL mean. This increase in female hiring for the older subsample in fact reinforces our results, as it suggests that employers may substitute older women for leave-likely women of child-bearing age.

5.3. Heterogeneity in labor market outcomes by industry

Finally, I seek to understand in which industries the labor-market impact of CA-PFL is most concentrated. For the subset of young workers, DDDs for each industry in the sample are presented in Table 8 for employment (Column 1), hires (Column 2) and separations (Column 3). Only the coefficient on the DDD interaction term of interest is reported. The publicly available LEHD data provide very coarse industry categories, but the results are interesting.

There is a negative and statistically significant decrease in female hiring in the retail, management, and administrative support industries. Separations decrease for women in retail and management as well, and also for women in mining/oil and gas, the arts, and the accommodation and food services industries. These crude results indicate that leave-taking patterns differ by industry[30] and are therefore differentially disruptive to employers depending on the sector. Interestingly, in some industries (educational services, healthcare, mining, and agriculture), female hires at large versus small firms increase during this period, and female separations at large versus small firms also increase in some of these same industries during this period (with the exception of educational services, where there is no statistical significance, and with the addition of construction). Overall employment effects in Column 1 vary depending on the industry in ways that affect the relative magnitudes of the hiring and separations results in Columns 2 and 3.

[30] Relatedly, Bartel et al. (2015) and that paternity leave increases as a result of CA-PFL are concentrated in occupations with a high share of female workers

It is not obvious why, in some industries, our results move in the opposite direction of our baseline results. It is worth noting that construction and mining have experienced changes related to the displacement of primarily male, low-skill workers during and after the Great Recession (Elsby and Hobijin 2010) that may have impacted the relative demand for male/female employees in a way completely unrelated to the passage of CA-PFL. The other industries (education and healthcare) are where job growth, especially for women, has and will be concentrated in the post-Recession epoch,[31] and it is possible this too is related to these results.

For the industries in which the results are directionally consistent with paid leave, reducing both hiring and separations (management and retail trade), I cannot comment on whether the hiring and separation results are distinct or related phenomenon. That is, one possibility is that a fraction of the workforces in these industries is traditionally women of child-bearing age. As a result of the provision of job-protected paid leave with CA-PFL, when these young women become pregnant, rather than leaving the workforce (which they would have done had leave not been available to them), many take advantage of CA-PFL and do not separate from their employers. Hence, there is a decrease in female separations for this subgroup. Employers then shift away from hiring women of child-bearing age because their young female employees are less likely to separate from their jobs post-CA-PFL.

An alternative possibility may be that these are two distinct phenomena. Now that job-protected paid leave is mandated, employers would like to reduce the share of their leave-likely employees. They cannot fire young women who are already employed and opt for paid leave because firing costs are too severe, but they actively choose to under-hire from this group to avoid having to bear the cost of future burdensome leaves. Future research should consider the extent to which the

[31] 2015 BLS projections suggest that healthcare and social assistance and educational services, will be the fastest-growing sectors in the next decade.

demand-side decision of employers to decrease female hiring in response to CA-PFL is a response to supply-side dynamics (decrease in female separations).

6. Conclusion

This paper contributes to the literature on parental leave in various ways. First, I build on the results of Rossin-Slater, Ruhm, and Waldfogel (2013) and establish that the increase in leave-taking they document as a result of CA-PFL is most pronounced for new mothers at large firms who have access to job protection through the FMLA. Although there are gains for small-firm employees as well (who have access to paid leave, but not job protection), these are more limited in scope. Also, gains for disadvantaged (low-educated, unmarried, and minority) new mothers who cannot afford to take unpaid FMLA leave exist only at large firms, where leave is protected.

Armed with the knowledge that CA-PFL increases leave-taking for large-firm female employees more than their small-firm counterparts, I compare employment outcomes at large versus small firms within California to large versus small firms outside of California to ascertain the labor-market implications of job-protected paid leave. This analysis builds on prior work that studies the impact of the extension of paid leave on female labor-force attachment and employment outcomes. Past authors have found relatively small medium- and long-term effects of leave policies on female employment (for example, Ruhm 1998; Waldfogel 1999; Baker and Milligan 2008; Han, Ruhm, and Jane Waldfogel 2009; and Baum and Ruhm 2016). I disentangle the overall employment impact into two distinct and important reactions to the provision of paid family leave in California. On the demand side, as Gruber (1992) and Summers (1989) caution with regard to group-specific mandates, employers actively substitute away from young women who are likely to take costly leave post-CA-PFL, and the percentage of new female hires falls by 1.1%. On the supply side, female employees who would have chosen to leave their jobs after giving birth instead

take job-protected paid leave, increasing female labor-force attachment, with the share of female separations decreasing by 0.07 percentage points (1.5%). The separations result appears slightly larger in magnitude, with overall employment increasing slightly for women at large versus small firms as a result of CA-PFL.

There are, of course, important limitations to the results presented. Although these findings are robust to the inclusion of a number of controls and placebo tests, data limitations prevent us from being able to ascertain (1) exactly how many employees work at each firm, and (2) the educational backgrounds of firm employees of different ages. As a result, this preliminary work is unable to explicitly take the hypothesis that low-skill women bear the brunt of the negative employment impact of job-protected leave to the data. It is also not possible to determine whether firms actively try to stay under the 50-employee cutoff as a result of CA-PFL. Such bunching would bias the results downward, suggesting that my findings provide a lower bound on the impact of CA-PFL on female labor-market outcomes. Further work should include comparisons to Rhode Island, where paid leave is job-protected for employees of all firm sizes. I also only consider whether women are in the labor force (employed, hired, or separated from their current employer). Recent work by Stearns (2016) considers other labor-market outcomes, cautioning that overly generous leaves may increase female employment but actually hinder career advancement, and Ruhm (1998) notes that longer paid-leave entitlements may have a negative long-term wage impact for female employees.

To formulate optimal leave policy, it will be important to understand the extent to which the two main employment results in this paper— (1) that employers substitute away from hiring leave-likely women post-CA-PFL and (2) that leave-likely women are more attached to a workplace that is more exible—are distinct or related. If employers choose to hire fewer young women because

the costs associated with parental leave (like finding a temporary worker to substitute for an employee on leave) are too high, policies such as those in the Norway, Sweden, Germany, and Finland, which force men to take leave after the birth of a newborn,[32] will be valuable not only for the impact they will have on children and new fathers, but also because such programs will make it harder for employers to substitute away from leave-likely women and instead hire men. If, however, the decision to hire fewer women is reactionary and the result of fewer young women choosing to separate from a workplace that is exible, = it is less clear that mandating leave for new fathers will meaningfully impact employers' hiring decisions

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[32] For a brief discussion of these programs, see Jackson (2015).

References

Acemoglu, D, and J. D. Angrist. 2001. "Consequences of Employment Protection? The Case of the Americans with Disabilities Act." *Journal of Political Economy* 109(5), 915-957.

Addati, L., N. Cassirer, and K. Gilchrist. 2014. *Maternity and Paternity at Work: Law and Practice Across the World*. International Labour Organization.

Appelbaum, E., and R. Milkman. 2011. *Leaves That Pay: Employer and Worker Experiences with Paid Family Leave in California.* Center for Economic and Policy Research.

Bana, S., K. Bedard, and M. Rossin-Slater. 2017. *The Impacts of Paid Family Leave Benefits: Regression Kink Evidence from California Administrative Data*. Working Paper No. 24438. The National Bureau of Economic Research.

Baker, M., and K. Milligan. 2008. "How Does Job-Protected Maternity Leave Affect Mothers' Employment?" *Journal of Labor Economics* 26(4), 655-692.

2015. "Maternity leave and children's cognitive and behavioral development." *Journal of Population Economics* 28(2), 373-391.

Bartel, A., M. Rossin-Slater, C. Ruhm, J. Stearns, and J. Waldfogel. 2015. *Paid Family Leave, Fathers' Leave-Taking, and Leave-Sharing in Dual-Earner Households.* Working Paper No. w21747. The National Bureau of Economic Research.

Barrera, J. 2017. *Job Killer Mandates New Leave of Absence*. California Chamber of Commerce Advocacy.

Baum, C. L. 2003. “The Effect of State Maternity Leave Legislation and the 1993 Family and Medical Leave Act on Employment and Wages.” *Labour Economics* 10(5), 573-596.

Baum, C. L., and C. J. Ruhm. 2016. The effects of paid family leave in California on labor market outcomes. *Journal of Policy Analysis and Management* 32(2), 224-245.

Bertrand, M., and S. Mullainathan. 2003. “Enjoying the quiet life? Corporate governance and managerial preferences.” *Journal of Political Economy* 111(5), 1043–1075.

Bravo, E. 2015. *Why Facebook CEO’s Paternity Leave is a Big Deal*. CNBC. November 24.

Byker, T. S. 2016. “Paid Parental Leave Laws in the United States: Does Short-Duration Leave Affect Women’s Labor-Force Attachment?” *American Economic Review* 106(5), 242-46.

Carneiro, P., K. Loken, and K. G. Salvanes. 2015. “A Flying Start: Maternity Leave Benefits and Long Run Outcomes of Children.” *Journal of Political Economy* 123(2), 365412.

Clinton, H. 2015. Paid Family and Medical Leave. Policy Statement.

Corte, R.L. 2017. *Inslee Signs Bill Guaranteeing Paid Family Leave in Washington*. Associated Press. July 5.

Das, T. and S. Polachek. 2015. “Unanticipated Effects of California’s Paid Leave Program.” *Contemporary Economic Policy* 33(4), 619-635.

DeLeire, T. 2000. “The Wage and Employment Effects of the Americans with Disabilities Act.” *Journal of Human Resources* 35(4), 693-715.

Donald, S. G., and K. Lang. 2007. “Inference with difference-in-differences and other panel data.” *The Review of Economics and Statistics* 89(2), 221-233.

Dustmann, C., and U. Schönberg. 2012. “Expansions in maternity leave coverage and children’s long-term outcomes.” *American Economic Journal: Applied Economics* 4(3), 190224.

Elsby, M. and B. Hobijin. 2010. *The Labor Market in the Great Recession*. Brookings Papers on Economic Activity.

Employment Development Department State of California. *Paid Family Leave* [Factsheet].

Engeman, C. 2012. *Ten Years of the California Paid Family Leave Program: Strengthening Commitment to Work, Affirming Commitment to Family*. Policy Brief.

Family and Medical Leave Act. 29 USC 2601.

National Partnership for Women and Families. 2015. *Family Medical Leave Act at 22: 200 Million Reasons to Celebrate and Move Forward.*

FindLaw. *FMLA Eligibility*.

Fass, S. 2009. *Paid Leave in the States: A Critical Support for Low-wage Workers and Their Families.* Technical Report. National Center for Children in Poverty.

Garthwaite, C., Gross, T., and Otidid, M. J. 2014. “Public health insurance, labor supply, and employment lock.” *The Quarterly Journal of Economics* 129(2), 653-696.

Goodman, M. 2015. *Can you be Fired for Taking Family Leave?* American Civil Liberties Union of Northern California.

Gruber, J. 1992. *The efficiency of a group-specific mandated benefit: Evidence from health insurance benefits for maternity*. Working Paper No. w4157. The National Bureau of Economic Research.

Gruber, J. 1994. "State-mandated benefits and employer-provided health insurance." *Journal of Public Economics* 55(3), 433-464.

Gruber, J. 1994. "The Incidence of Mandated Maternity Benefits." *The American Economic Review* 84(3), 622-641.

Han, W-J., and J. Waldfogel. 2003. "Parental Leave: The Impact of Recent Legislation on Parents' Leave Taking." *Demography* 40(1), 191-200.

Han, W-J., C. Ruhm, and Jane Waldfogel. 2009. "Parental Leave Policies and Parents' Employment and Leave-Taking." *Journal of Policy Analysis Management* 28(1), 29-54.

Houser, L., and T. Vartanian. 2012. *Pay Matters: The Positive Economic Impacts of Paid Family Leave for Families, Businesses and the Public.* New Brunswick, NJ: Rutgers Center for Women and Work.

Jackson, G. 2015. "Force Men to Take Paternity Leave. It will Make the World a Better Place." *The Guardian*. April 9.

Jolls, Christine. "Accommodation mandates." *Stanford Law Review*. 2000: 223-306.

Jaumotte, F. 2003. "Labour Force Participation of Women: Empirical Evidence on the Role of Policy and Other Determinants in OECD countries." *OECD Economic Studie*s 2, 51-108.

Klerman J. A., K. Daley, and A. Pozniak. 2012. *Family and medical leave in 2012: Technical report.* Abt Associates Inc. Prepared for the U.S. Department of Labor.

U.S. Census Bureau. 2015. *Local Employment Dynamics.*

International Labour Organization. 2014. *Maternity and Paternity at Work: Law and Practice Around the World.*

Miller, Clare Cain. 2017. "Family-Friendly Laws are Being Passed, but Not by Trump's Team." *The New York Times*. July 3.

Ondrich, J.C., K. Spiess, and Q. Yang. 1996. "Barefoot and in a German kitchen: Federal Parental Leave and Benefit Policy and the Return to Work after Childbirth in Germany." *Journal of Population Economics* 9(3), 247-266.

Pelletier, A. 2007. "The Family Medical Leave Act of 1993-Why Does Parental Leave in the United States Fall So Far Behind Europe?" *Gonzaga Law Review* 42(3).

Employment Development Department. State of California. *PFL Program Statistics*.

Price, J. 2008. "Parent-Child Quality Time Does Birth Order Matter?" *Journal of Human Resources* 43(1), 240-265.

Reagan, K. 2013. *The 'Illusion' of Family Leave in New York*.

Ross, A. 2013. *It Pays to Leave: The Impact of Paid Family Leave on Mothers in California.* Unpublished manuscript. Harvard University.

Rossin-Slater, M. 2017. *Maternity and Family Leave Policy*. Working Paper No. w23609. The National Bureau of Economic Research.

Rossin-Slater, M., C. J. Ruhm, and J. Waldfogel. 2013. "The Effects of California's Paid Family Leave Program on Mothers' Leave-Taking and Subsequent Labor Market Outcomes." *Journal of Policy Analysis and Management* 32(2), 224-245.

Rossin, M. 2011. "The effects of maternity leave on children's birth and infant health outcomes in the United States." *Journal of Health Economics* 30(2), 221-239.

Ruggles, S., K. Genadek, R. Goeken, J. Grover, and M. Sobek. 2015. *Integrated Public Use Microdata Series: Version 6.0* [dataset]. Minneapolis: University of Minnesota.

Ruhm, C. J. 2000. "Parental Leave and Child Health." *Journal of Health Economics* 19(6), 931-960.

1998. "The Economic Consequences of Parental Leave Mandates: Lessons from Europe." *The Quarterly Journal of Economics* 113(1), 285-317.

Stearns, J. 2015. "The effects of paid maternity leave: Evidence from Temporary Disability Insurance." *Journal of Health Economics* 43, 85-102.

Employment Development Department. State of California. *SDI Contribution Rates*.

A Better Balance. *The Need for Paid Family Leave*.

Summers, L.H. 1989. Some Simple Economics of Mandated Benefits. *The American Economic Review* 79(2), 177-183.

Sullivan, S., and R. Costa. 2016. “Donald Trump Unveils Child-Care Policy Influenced by Ivanka Trump.” *The Washington Post*. September 13.

Waldfogel, J. 2001. “Family and Medical Leave: Evidence from the 2000 Surveys.” *Monthly Labor Review* 124(9), 17-23.

1999. “The Impact of the Family and Medical Leave Act.” *Journal of Policy Analysis and Management* 18(2), 282-302.

Wojcicki, S. 2014. “Paid Maternity Leave is Good for Business.” *The Wall Street Journal*. December 16.

Table 1. Descriptive statistics for selected analysis variables

		Treatment: CA mothers of youngest children aged <1		Comparison: CA mothers of youngest children aged 5–17	
Panel A: All firms	Full sample	Pre	Post	Pre	Post
On maternity leave last week	1.0	5.3	11.4	0.2	0.1
On family leave last week	2.7	8.0	14.6	1.5	1.7
Mother is non-Hispanic white	42.5	43.7	41.2	45.1	39.9
Mother is black	6.7	5.3	5.7	7.0	6.7
Mother is Hispanic	35.6	35.6	38.3	32.8	38.2
Mother is married	70.9	77.3	76.2	70.0	70.3
Mother ed: <HS	17.2	14.1	13.4	17.7	17.6
Mother ed: HS degree	22.8	25.9	17.8	24.3	21.6
Mother ed: Some college	31.3	30.3	30.2	32.6	30.2
Mother ed: College or more	28.7	29.7	38.6	25.4	30.7
Sample size	15,319	682	825	6,222	7,590
Panel B: Large firms					
On maternity leave last week	1.3	7.1	13.5	0.2	0.2
On family leave last week	3.0	10.1	17.6	1.5	1.7
Mother is non-Hispanic white	43.1	43.6	45.1	45.3	40.6
Mother is black	8.7	6.4	7.0	9.5	8.3
Mother is Hispanic	32.5	32.5	34.8	29.9	34.8
Mother is married	70.8	78.8	77.6	69.9	69.8
Mother born in United States	65.0	68.9	72.3	66.5	62.0
Mother ed: <HS	13.5	8.9	10.5	14.2	13.7
Mother ed: HS degree	22.0	27.0	16.5	24.0	20.2
Mother ed: Some college	31.3	30.3	29.0	32.7	30.3
Mother ed: College or more	33.2	33.8	44.0	29.1	35.8
Sample size	8,643	397	483	3554	4,209
Panel C: Small firms					
On maternity leave last week	0.6	2.4	7.0	0.1	0.1
On family leave last week	2.4	5.3	9.5	1.6	2.1
Mother is non-Hispanic white	43.4	48.0	35.5	46.5	41.0
Mother is black	3.8	4.1	3.4	3.4	4.2
Mother is Hispanic	38.1	40.4	43.1	34.6	40.6
Mother is married	71.7	74.1	73.3	70.8	72.0
Mother ed: <HS	21.9	24.1	18.1	21.7	22.2
Mother ed: HS degree	24.0	23.7	19.4	26.1	22.6
Mother ed: Some college	31.6	29.7	30.9	32.4	31.1
Mother ed: College or more	22.6	22.5	31.5	19.8	24.2

Sample size	4,836	207	251	1,911	2,467

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), the data comes from the 1999-2010 March CPS Surveys. All statistics are weighted by the March CPS Supplement person weights. The pre period refers to 1999-2004; the post period refers to 2005-2010. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who live in California and reported working any usual hours in the last year. Maternity leave includes mothers that respond that they were absent from work last week due to maternity leave; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees. For brevity, I provide summary statistics for key variables, not the whole group of variables considered by Rossin-Slater, Ruhm, and Waldfogel (2013). A complete replication of their descriptive statistics for the whole sample, and then a split by firm size, is available in the Appendix in Tables A1, Panels A-C.

Table 2: Estimated effects of California paid family leave on leave-taking

Panel A: All firms	Maternity leave	Family leave	Other leave	Any leave
Estimated PFL effect	5.98***	6.09***	-0.41	5.57***
	(1.15)	(1.07)	(0.84)	(1.40)
Panel B: Large firms				
Estimated PFL effect	6.41***	7.23***	0.10	6.51***
	(1.70)	(1.85)	(0.78)	(1.82)
Panel C: Small firms				
Estimated PFL effect	4.57**	3.72*	-1.99	2.58
	(1.51)	(2.00)	(2.13)	(2.10)

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), Table 2: Each coefficient is from a separate regression, with standard errors in parentheses. The data comes from the 1999 to 2010 March CPS Surveys. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who live in California and reported working any usual hours in the last year. The treatment group includes women with a youngest child aged <1 year, while the comparison group includes women with a youngest child aged 5 to 17 years. All regressions include controls for age buckets (<20, 20 to 29, 30 to 39, 40 to 49, 50 to 59, 60+), indicators for race and ethnicity (non-Hispanic white, black, Hispanic, other), indicators for marital status (married, divorced, separated, widowed, never married), an indicator for being born in the United States, indicators for education categories (< HS, HS, some college, college or more), and indicators for single years of the age of the youngest child. All regressions are weighted by March CPS person weights. The estimated PFL effect is computed using the Donald and Lang (2007) two-step method since there is only one state in the sample. In the first step, each outcome is regressed on the full set of controls, survey year dummies, and the year dummies interacted with treatment status, with no constant. In the second step, data are collapsed into survey-year cells and the coefficient on the interaction between treatment status and year is regressed on an indicator for post-2004. The coefficient and corresponding standard error on the post-2004 indicator is reported here, and statistical significance is determined using a t-distribution with 10 degrees of freedom. Maternity leave indicates mothers who were employed and respond that they were absent from the job due to maternity leave last week; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons. Other leave includes all reasons for job absences except for maternity leave. Any leave includes mothers employed but absent from work for any reason. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees.

Table 3. Subgroup estimates of leave use

Panel A: All firms	N	Pre-PFL mean (%)	Estimated PFL effect	
			Coef.	(SE)
Mothers with high school degree or less	6,448	2.2	4.69**	(1.77)
Mothers with some college	4,666	5.5	6.32**	(2.55)
Mothers with college degree or more	4,155	9.3	5.11	(3.65)
Unmarried mothers	4,438	2.2	6.38***	(1.39)
Married mothers	10,831	6.2	6.04***	(1.41)
Non-Hispanic white mothers	5,976	6.9	4.16	(2.75)
Black mothers	880	2.1	9.87*	(5.15)
Hispanic mothers	6,434	3.9	5.06***	(1.33)
Panel B: Large firms				
Mothers with high school degree or less	3,190	2.2	7.66**	(2.93)
Mothers with some college	2,667	7.8	5.95*	(3.19)
Mothers with college degree or more	2,747	11.6	5.53	(3.92)
Unmarried mothers	2,505	1.7	8.59**	(2.92)
Married mothers	6,109	8.5	5.79**	(1.98)
Non-Hispanic white mothers	3,496	9.6	4.53	(3.36)
Black mothers	639	2.8	7.83	(5.76)
Hispanic mothers	3,312	5.4	5.28*	(2.48)
Panel C: Small firms				
Mothers with high school degree or less	2,329	1.7	1.06	(1.80)
Mothers with some college	1,461	1.5	6.24	(3.95)
Mothers with college degree or more	1,030	4.9	5.90	(6.32)
Unmarried mothers	1,372	2.0	1.68	(3.38)
Married mothers	3,448	2.5	6.78**	(2.35)
Non-Hispanic white mothers	1,885	2.3	2.87	(3.64)
Black mothers	163	0.0	−3.29	(4.66)
Hispanic mothers	2,169	1.9	2.87	(2.34)

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), Table 6: Each coefficient is from a separate regression, with standard errors in parentheses. The data comes from the 1999 to 2010 March CPS Surveys. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who live in California and reported working any usual hours in the last year within each of the subgroups listed above. The treatment group includes women with a youngest child aged <1 year, while the comparison group includes women with a youngest child aged 5 to 17 years. Depending on the relevant subgroup, regressions include controls for age buckets (<20, 20 to 29, 30 to 39, 40 to 49, 50 to 59, 60+), indicators for race and ethnicity (non-Hispanic white, black, Hispanic, other), indicators for marital status (married, divorced, separated, widowed, never married), an indicator for being born in the United States, indicators for education categories (< HS, HS, some

college, college or more), and indicators for single years of the age of the youngest child. All regressions are weighted by March CPS person weights. The estimated PFL effect is computed using the Donald and Lang (2007) two-step method since there is only one state in the sample. In the first step, each outcome is regressed on the full set of controls, survey year dummies, and the year dummies interacted with treatment status, with no constant. In the second step, data are collapsed into survey-year cells and the coefficient on the interaction between treatment status and year is regressed on an indicator for post-2004. The coefficient and corresponding standard error on the post-2004 indicator is reported here, and statistical significance is determined using a t-distribution with 10 degrees of freedom. Maternity leave indicates mothers who were employed and respond that they were absent from the job due to maternity leave last week. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees.

Table 4. Descriptive statistics for Census LEHD sample

	% Female Employees		% Female Hires		% Female Separations	
	Pre	Post	Pre	Post	Pre	Post
California						
All firms	43.2	42.9	4.1	4.6	4.2	4.6
Younger workers	44.2	44.1	5.6	6.1	5.0	5.3
Older workers	41.7	41.2	2.0	2.4	3.1	3.4
Small Firms	42.9	42.4	4.0	4.5	4.3	4.6
Younger workers	44.2	43.5	5.4	5.9	5.1	5.4
Older workers	41.1	40.9	2.0	2.4	3.2	3.4
Large Firms	42.3	41.9	4.0	4.4	4.0	4.5
Younger workers	43.4	43.2	5.3	5.8	4.8	5.2
Older workers	40.6	40.2	2.0	2.3	2.9	3.4
Other states (excluding NJ)						
All firms	43.5	43.6	3.9	3.9	4.1	4.3
Younger workers	44.2	44.6	5.3	5.3	4.9	5.1
Older workers	42.4	42.0	1.7	1.8	2.9	3.0
Small Firms	43.6	43.4	3.7	3.8	4.1	4.2
Younger workers	44.4	44.3	5.0	5.1	4.9	5.1
Older workers	42.4	42.0	1.7	1.8	2.8	2.9
Large Firms	43.0	43.4	3.8	3.9	4.0	4.3
Younger workers	43.7	44.4	5.1	5.3	4.7	5.1
Older workers	42.0	41.7	1.7	1.8	2.8	2.9

Note: Data are for employed civilians under the age of 45 from the 2000-2010 Census LEHD database. The pre period refers 2000-2004; the post period refers to 2005-2010. All estimates are reported as percentages. The large firm subsample includes women who report working at firms with more than 50-249 employees; the small firm subsample includes women who report working at firms with 25-49 employees.

Table 5. Effect of Job-Protected Paid Leave in California on Employment, Hires, and Separations Relative to Other States (DDD)

	Employment	Hires	Separations
	(1)	(2)	(3)
isCA x Large x Post	0.652***	-0.0609**	-0.0706***
	(0.187)	(0.0249)	(0.0252)
Mean of dependent variable	44.23	5.17	4.85
Fixed Effects: Year-Quarter, State, Size, Age group, Industry	Yes	Yes	Yes
Observations	536,150	536,150	536,150
R-squared	0.880	0.644	0.587

Note: Data are for employed civilians under the age of 45 from the 2000-2010 Census LEHD database. This table reports the results of OLS quarterly regressions relating workforce composition to CA-PFL. Results are reported from a DDD specification comparing female employees of childbearing age (under 45 years) at small firms and large firms in California to small and large firms outside of California. The dependent variable is percentage of female employees, percentage of female hires, and percentage of female separations, depending on specification. Robust standard errors are clustered at the state level. isCA is an indicator that takes a value of 1 if the firm is in California, Post is an indicator variable that takes value 1 after the passage of paid leave in Q3 2002. Large is an indicator that takes a value of 1 if the firm is large (50-249 employees).

Table 6. Effect of Job-Protected Paid Leave in California Relative to Other States (DDD), Placebo Size Test

	Small versus smaller firms			Large versus larger firms		
	Employment	Hires	Separations	Employment	Hires	Separations
	(1)	(2)	(3)	(4)	(5)	(6)
isCA x Large x Post	0.375*	0.180***	0.194***	-0.508*	0.202***	0.0785**
	(0.191)	(0.0406)	(0.0338)	(0.282)	(0.0467)	(0.0388)
Mean of dependent variable	44.25	5.18	5.04	44.00	4.93	4.60
Fixed Effects: Year-Quarter, State, Size, Age group, Industry	Yes	Yes	Yes	Yes	Yes	Yes
Observations	537,424	537,424	537,424	519,893	519,893	519,893
R-squared	0.887	0.673	0.636	0.818	0.566	0.505

Table 7. Effect of Job-Protected Paid Leave in California on Employment, Hires, and Separations Relative to Other States (DDD), Placebo Age Test

	Employment	Hires	Separations
	(1)	(2)	(3)
isCA x Large x Post	0.283	0.0305**	-0.0561
	(0.255)	(0.0135)	(0.0357)
Mean of dependent variable	42.18	1.69	2.90
Fixed Effects: Year-Quarter, State, Size, Age group, Industry	Yes	Yes	Yes
Observations	325,887	325,887	325,887
R-squared	0.905	0.620	0.551

Note: Data are for employed civilians from the 2000-2010 Census LEHD database. This table reports the results of OLS quarterly regressions relating workforce composition to CA-PFL. In Table 6 (Table 7), results are reported from a DDD specification comparing female employees of childbearing age, under 45 years (over childbearing age, 45+) at small firms and large firms in California to small and large firms outside of California. The dependent variable is percentage of female employees, percentage of female hires, and percentage of female separations, depending on specification. Robust standard errors are clustered at the state level. isCA is an indicator that takes a value of 1 if the firm is in California, Post is an indicator variable that takes value 1 after the passage of paid leave in Q3 2002. In Table 6, Large is an indicator that takes a value of 1 if the firm is large (2049 employees in the small placebo test, 250-499 employees in the large placebo test). In Table 7, Large is an indicator that takes a value of 1 if the firm is large (50-249 employees). I include year-quarter, age group, industry, geography, and firm-size fixed effects, as well as all possible pairwise interactions of these fixed effects. The use of these fixed effects makes it unnecessary to control explicitly for isCA, Post, Large, or any pairwise interaction of these variables.

Table 8. Effect of Job-Protected Paid Leave in California Relative to Other States (DDD), by Industry

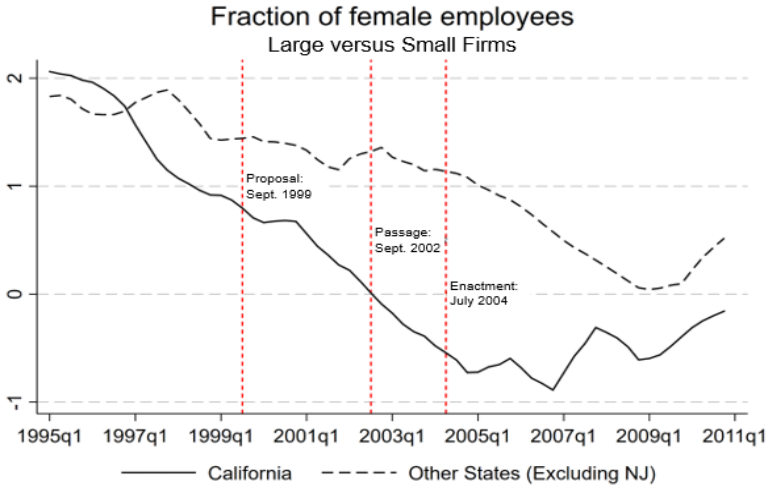
Industry Subsample	Employment (1)	Hires (2)	Separations (3)
Agriculture, Forestry, Fishing and Hunting	2.378**	0.217*	0.573***
	(0.964)	(0.128)	(0.166)
Mining, Quarrying, and Oil and Gas Extraction	0.759	0.274***	-0.276***
	(0.812)	(0.069)	(0.067)
Utilities	7.430***	0.032	-0.000002
	(2.410)	(0.185)	(0.234)
Construction	-0.270	-0.0124	0.0976**
	(0.396)	(0.049)	(0.037)
Manufacturing	0.381	0.0705	0.153
	(0.614)	(0.081)	(0.093)
Wholesale Trade	0.671	0.0396	-0.0284
	(0.599)	(0.070)	(0.099)
Retail Trade	0.690**	-0.145***	-0.0992*
	(0.309)	(0.050)	(0.055)
Transportation and Warehousing	-0.644	-0.146	-0.0834
	(0.469)	(0.111)	(0.100)
Information	0.999	0.108	-0.199
	(0.622)	(0.157)	(0.119)
Finance and Insurance	0.888*	-0.109	0.0822
	(0.520)	(0.128)	(0.136)
Real Estate and Rental and Leasing	0.277	0.177	0.216
	(1.154)	(0.154)	(0.190)
Professional, Scientific, and Technical Services	0.492	-0.0912	-0.0512

	(0.573)	(0.144)	(0.100)
Management of Companies and Enterprises	-2.382**	-1.507***	-1.161***
	(1.070)	(0.190)	(0.289)
Administrative and Support and Waste Management and Remediation Services	-0.0731	-0.317**	-0.0141
	(0.926)	(0.135)	(0.169)
Educational Services	2.750***	0.956***	0.259
	(1.013)	(0.247)	(0.243)
Health Care and Social Assistance	1.491***	0.203***	0.203**
	(0.487)	(0.063)	(0.082)
Arts, Entertainment, and Recreation	-2.233***	-0.0576	-0.355**
	(0.735)	(0.123)	(0.174)
Accommodation and Food Services	-0.265	-0.0326	-0.0743*
	(0.175)	(0.042)	(0.040)
Other Services (except Public Administration)	-1.525***	-0.213**	-0.144
	(0.521)	(0.099)	(0.096)

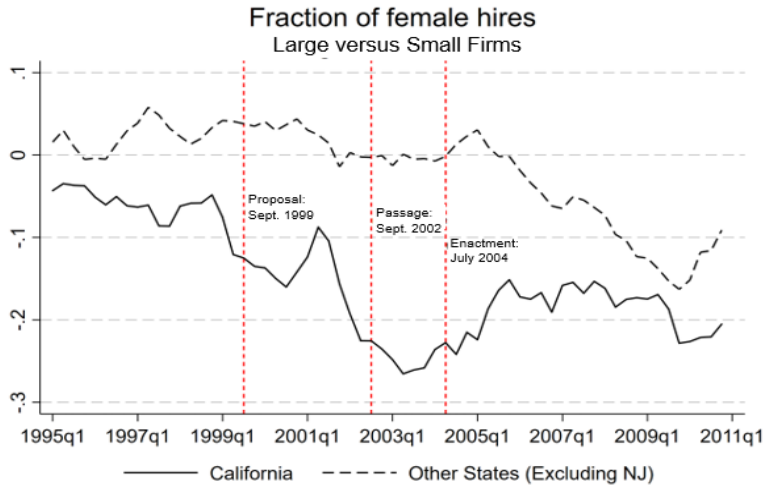
Note: Data are for employed civilians under the age of 45 from the 2000-2010 Census LEHD database. This table reports the results of OLS quarterly regressions relating workforce composition to CA-PFL. Results are reported from a DDD specification comparing female employees of childbearing age (19-45) at small firms (2049 employees) and large firms (50-249 employees) in CA to small and large firms outside of CA within each industry listed above. The dependent variable is percentage of female employees (Col. 1), hires (Col. 2), and separations (Col. 3). Robust standard errors are clustered at the state level. The regression specification follows that of Table 5 (except that it is estimated for a particular industry subgroup). I include year-quarter, age group, geography, and firm size fixed effects, as well as all possible pairwise interactions of these fixed effects. The use of these fixed effects makes it unnecessary to control explicitly for isCA, Post, Large, or any pairwise interaction of these variables. Only the main coefficient of interest, isCA x Large x Post, is reported.

Figure 1. Pretrends: Female Employment at Small and Large Firms in California versus Control States

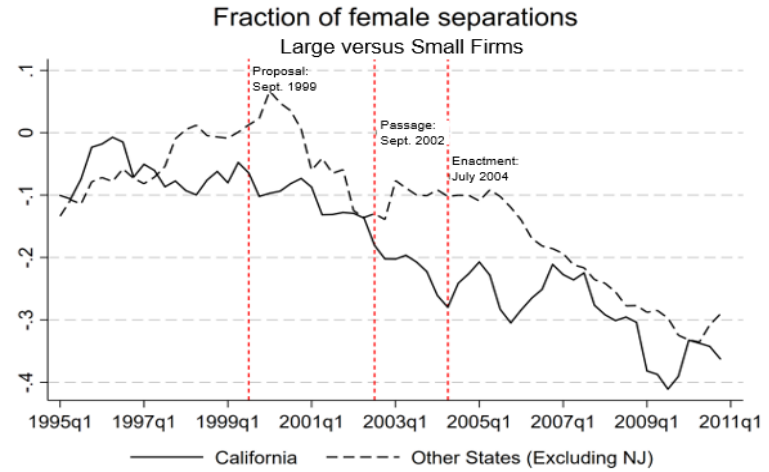
Panel A: Employment



Panel B: Hires



Panel C: Separations



Note: Data are for employed civilians under the age of 45 from the 1995–2010 Census LEHD database. These figures compare fraction of female employees (Panel A), hires (Panel B), and separations (Panel C) at small firms (20–49 employees) and large firms (50–249 employees) in California to small and large firms in non-treatment states. In each quarter, I report the difference of the dependent variable for large and small firms separately within and outside California, across all industries (and all geographies, for the control states). For example, for female employment, I plot:

$$ShareFemEmp\{z = Large\}\{s = CA\} - ShareFemEmp\{z = Small\}\{s = CA\}$$

and

$$ShareFemEmp\{z = Large\}\{s = NoCA\} - ShareFemEmp\{z = Small\}\{s = NoCA\}$$

Table A1. Descriptive statistics for selected analysis variables

Panel A: All firms	Full sample	Treatment: CA mothers of youngest children aged<1		Comparison: CA mothers of youngest children aged 5–17	
		Pre	Post	Pre	Post
On maternity leave last week	1.0	5.3	11.4	0.2	0.1
On family leave last week	2.7	8.0	14.6	1.5	1.7
On other leave last week	2.4	4.2	4.1	2.2	2.2
On any leave last week	3.5	9.5	15.5	2.3	2.4
With job and at work last week	84.2	59.5	57.9	86.5	88.1
Not employed last week	12.2	31.0	26.1	11.2	9.4
Age of youngest child (0 means < 1)	9.5	0.0	0.0	10.5	10.9
Mother age: <20	0.3	3.3	2.7	0.0	0.0
Mother age: 20 to 39	10.3	45.7	43.6	6.7	5.6
Mother age: 30 to 39	35.2	43.3	48.7	36.4	31.3
Mother age: 40 to 49	43.6	7.6	5.0	47.4	48.9
Mother age: 50 to 59	10.2	0.0	0.1	9.2	13.8
Mother is non-Hispanic white	42.5	43.7	41.2	45.1	39.9
Mother is black	6.7	5.3	5.7	7.0	6.7
Mother is Hispanic	35.6	35.6	38.3	32.8	38.2
Mother is other race	15.1	15.4	14.8	15.0	15.2
Mother is married	70.9	77.3	76.2	70.0	70.3
Mother is separated	4.9	2.3	2.3	5.0	5.3
Mother is divorced	11.9	2.6	2.2	13.8	12.3
Mother is widowed	12.4	17.8	19.3	11.1	12.1
Mother born in United States	60.8	67.5	68.7	62.2	57.6
Mother ed: <HS	17.2	14.1	13.4	17.7	17.6
Mother ed: HS degree	22.8	25.9	17.8	24.3	21.6
Mother ed: Some college	31.3	30.3	30.2	32.6	30.2
Mother ed: College or more	28.7	29.7	38.6	25.4	30.7
Sample size	15319	682	825	6222	7590

Panel B: Large firms					
On maternity leave last week	1.3	7.1	13.5	0.2	0.2
On family leave last week	3.0	10.1	17.6	1.5	1.7
On other leave last week	2.6	4.5	4.9	2.3	2.3
On any leave last week	3.9	11.6	18.4	2.4	2.5
With job and at work last week	84.3	58.5	56.1	87.1	88.5
Not employed last week	11.6	29.9	24.7	10.4	8.8
Age of youngest child (0 means < 1)	9.6	0.0	0.0	10.6	10.9
Mother age: <20	0.3	3.3	2.3	0.1	0.0
Mother age: 20 to 39	10.3	45.3	43.2	6.6	5.2
Mother age: 30 to 39	34.5	44.2	50.3	35.5	30.0
Mother age: 40 to 49	44.2	7.2	4.1	48.2	50.1
Mother age: 50 to 59	10.4	0.0	0.1	9.4	14.3
Mother is non-Hispanic white	43.1	43.6	45.1	45.3	40.6
Mother is black	8.7	6.4	7.0	9.5	8.3
Mother is Hispanic	32.5	32.5	34.8	29.9	34.8
Mother is other race	15.7	17.5	13.1	15.3	16.2

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Panel B: Large firms (continued)	Full sample	Treatment: CA mothers of youngest children aged<1 Pre	Treatment: CA mothers of youngest children aged<1 Post	Treatment: CA mothers of youngest children aged 5–17 Pre	Treatment: CA mothers of youngest children aged 5–17 Post
Mother is married	70.8	78.8	77.6	69.9	69.8
Mother is separated	4.6	2.1	2.4	4.9	4.8
Mother is divorced	12.2	3.9	1.2	14.1	12.8
Mother is widowed	12.4	15.2	18.8	11.2	12.5
Mother born in United States	65.0	68.9	72.3	66.5	62.0
Mother ed: <HS	13.5	8.9	10.5	14.2	13.7
Mother ed: HS degree	22.0	27.0	16.5	24.0	20.2
Mother ed: Some college	31.3	30.3	29.0	32.7	30.3
Mother ed: College or more	33.2	33.8	44.0	29.1	35.8
Sample size	8643	397	483	3554	4209
Panel C: Small Firms					
On maternity leave last week	0.6	2.4	7.0	0.1	0.1
On family leave last week	2.4	5.3	9.5	1.6	2.1
On other leave last week	2.6	4.8	3.2	2.3	2.5
On any leave last week	3.1	7.2	10.2	2.4	2.6
With job and at work last week	84.4	59.6	61.6	86.0	88.
Not employed last week	12.5	33.1	28.2	11.6	9.4
Age of youngest child (0 means <1)	9.6	0.0	0.0	10.4	10.9
Mother age: <20	0.3	3.2	2.6	0.0	0.0
Mother age: 20 to 39	9.7	44.3	44.5	6.6	5.1
Mother age: 30 to 39	35.7	42.7	45.6	38.0	31.7
Mother age: 40 to 49	43.2	9.8	7.3	45.3	48.9
Mother age: 50 to 59	10.8	0.0	0.0	9.7	14.1
Mother is non-Hispanic white	43.4	48.0	35.5	46.5	41.0
Mother is black	3.8	4.1	3.4	3.4	4.2
Mother is Hispanic	38.1	40.4	43.1	34.6	40.6
Mother is other race	14.6	7.6	18.0	15.5	14.2
Mother is married	71.7	74.1	73.3	70.8	72.0

Mother is separated	5.2	3.3	2.4	5.0	5.9
Mother is divorced	11.3	0.3	4.1	13.0	11.7
Mother is widowed	11.8	22.3	20.3	11.2	10.3
Mother born in United States	55.7	64.3	63.9	57.3	52.3
Mother ed: <HS	21.9	24.1	18.1	21.7	22.2
Mother ed: HS degree	24.0	23.7	19.4	26.1	22.6
Mother ed: Some college	31.6	29.7	30.9	32.4	31.1
Mother ed: College or more	22.6	22.5	31.5	19.8	24.2
Sample size	4836	207	251	1911	2467

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), Table 1 the data comes from the 1999-2010 March CPS Surveys. All statistics are weighted by the March CPS Supplement person weights. The pre period refers to 1999-2004; the post period refers to 2005-2010. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who live in California and reported working any usual hours in the last year. Maternity leave includes mothers that respond that they were absent from work last week due to maternity leave; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons. Other leave includes all reasons for job absences except for maternity leave. Any leave includes mothers employed but absent from work for any reason. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees.

Table A2. Estimated effects of California paid family leave on leave-taking: Rossin-Slater, Ruhm, and Waldfogel (2013) Replication

	Maternity leave	Family leave	Other leave	Any leave
Treatment-on-the-treated (TOT)				
Estimated PFL effect	5.98***	6.09***	−0.41	5.57***
	(1.15)	(1.07)	(0.84)	(1.40)
Intent-to-treat (ITT)				
Estimated PFL effect	3.39***	3.50***	−0.19	3.21***
	(0.70)	(0.70)	(0.55)	(0.92)
Implied TOT from ITT estimate	5.74	5.92	−0.32	5.43

Note: This Table replicates Rossin-Slater, Ruhm, and Waldfogel (2013), Table 2. Each coefficient is from a separate regression, with standard errors in parentheses. The data comes from the 1999 to 2010 March CPS Surveys. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who live in California. The TOT sample is further limited to only include those women who report any usual hours worked in the prior year. The treatment group includes women with a youngest child aged <1 year, while the comparison group includes women with a youngest child aged 5 to 17 years. All regressions include controls for age buckets (<20, 20 to 29, 30 to 39, 40 to 49, 50 to 59, 60+), indicators for race and ethnicity (non-Hispanic white, black, Hispanic, other), indicators for marital status (married, divorced, separated, widowed, never married), an indicator for being born in the United States, indicators for education categories (< HS, HS, some college, college or more), and indicators for single years of the age of the youngest child. All regressions are weighted by March CPS person weights. The estimated PFL effect is computed using the Donald and Lang (2007) two-step method since there is only one state in the sample. In the first step, each outcome is regressed on the full set of controls, survey year dummies, and the year dummies interacted with treatment status, with no constant. In the second step, data are collapsed into survey-year cells and the coefficient on the interaction between treatment status and year is regressed on an indicator for post-2004. The coefficient and corresponding standard error on the post-2004 indicator is reported here, and statistical significance is determined using a t distribution with 10 degrees of freedom. The implied TOT coefficient is calculated by dividing the ITT estimate by the pre-PFL treatment group rate (the share of respondents who report working any usual hours in the prior year, which is 59.1% here). Maternity leave indicates mothers who were employed and respond that they were absent from the job due to maternity leave last week; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons.

Other leave includes all reasons for job absences except for maternity leave. Any leave includes mothers employed but absent from work for any reason.

Table A3. Effects of California paid family leave on leave-taking with different comparison groups

Panel A: All firms	Maternity leave	Family leave	Other leave	Any leave
Comparison group 1: Mothers of youngest children aged 5 to 17	5.98***	6.09***	-0.41	5.57***
	(1.15)	(1.07)	(0.84)	(1.40)
Comparison group 2: Women with no children	6.01***	6.57***	-0.13	5.87***
	(1.17)	(1.08)	(0.90)	(1.52)
Comparison group 3: Mothers of youngest children aged <1 year in FL, NY, TX	4.66**	4.77**	-0.56	4.09*
	(1.56)	(1.77)	(1.31)	(2.19)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA	3.20**	3.74***	-0.36	2.84*
	(1.22)	(1.16)	(0.92)	(1.55)
Panel B: Large firms				
Comparison group 1: Mothers of youngest children aged 5 to 17	6.41***	7.23***	0.10	6.51***
	(1.70)	(1.85)	(0.78)	(1.82)
Comparison group 2: Women with no children	6.60***	7.68***	0.34	6.94***
	(1.68)	(1.84)	(0.87)	(1.98)
Comparison group 3: Mothers of youngest children aged <1 year in FL, NY, TX	4.91**	5.39*	-0.25	4.67*
	(2.15)	(2.60)	(1.54)	(2.46)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA	3.17*	4.18*	0.08	3.24
	(1.60)	(1.93)	(0.89)	(1.88)

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Panel C: Small firms				
Comparison group 1: Mothers of youngest children aged 5 to 17	4.57**	3.72*	-1.99	2.58
	(1.51)	(2.00)	(2.13)	(2.10)
Comparison group 2: Women with no children	4.57***	4.58**	-1.58	3.00
	(1.42)	(1.85)	(2.17)	(2.06)
Comparison group 3: Mothers of youngest children aged <1 year in FL, NY, TX	2.47	1.71	-2.07	0.40
	(2.45)	(3.17)	(2.45)	(2.98)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA	1.93	1.36	-2.08	-0.15
	(1.73)	(1.85)	(2.22)	(1.97)

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), Table 3: Each coefficient is from a separate regression, with standard errors in parentheses. The data comes from the 1999 to 2010 March CPS Surveys. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who reported working any usual hours in the last year. The treatment group includes women with a youngest child aged <1 year in CA, while the comparison groups include (1) women with a youngest child aged 5 to 17 years in CA; (2) women with no children in CA; (3) women with a youngest child aged <1 year in FL, NY, or TX; (4) women with a youngest child aged <1 year in all states except CA. All regressions include controls for age buckets (<20, 20 to 29, 30 to 39, 40 to 49, 50 to 59, 60+), indicators for race and ethnicity (non-Hispanic white, black, Hispanic, other), indicators for marital status (married, divorced, separated, widowed, never married), an indicator for being born in the United States, indicators for education categories (< HS, HS, some college, college or more), and indicators for single years of the age of the youngest child. All regressions are weighted by March CPS person weights. For comparison groups 1 and 2, the estimated PFL effect is computed using the Donald and Lang (2007) two-step method since there is only one state in the sample, following Table 2. For comparison groups 3 and 4, state fixed effects are included. Maternity leave indicates mothers who were employed and respond that they were absent from the job due to maternity leave last week; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons. Other leave includes all reasons for job absences except for maternity leave. Any leave includes mothers employed but absent from work for any reason. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees.

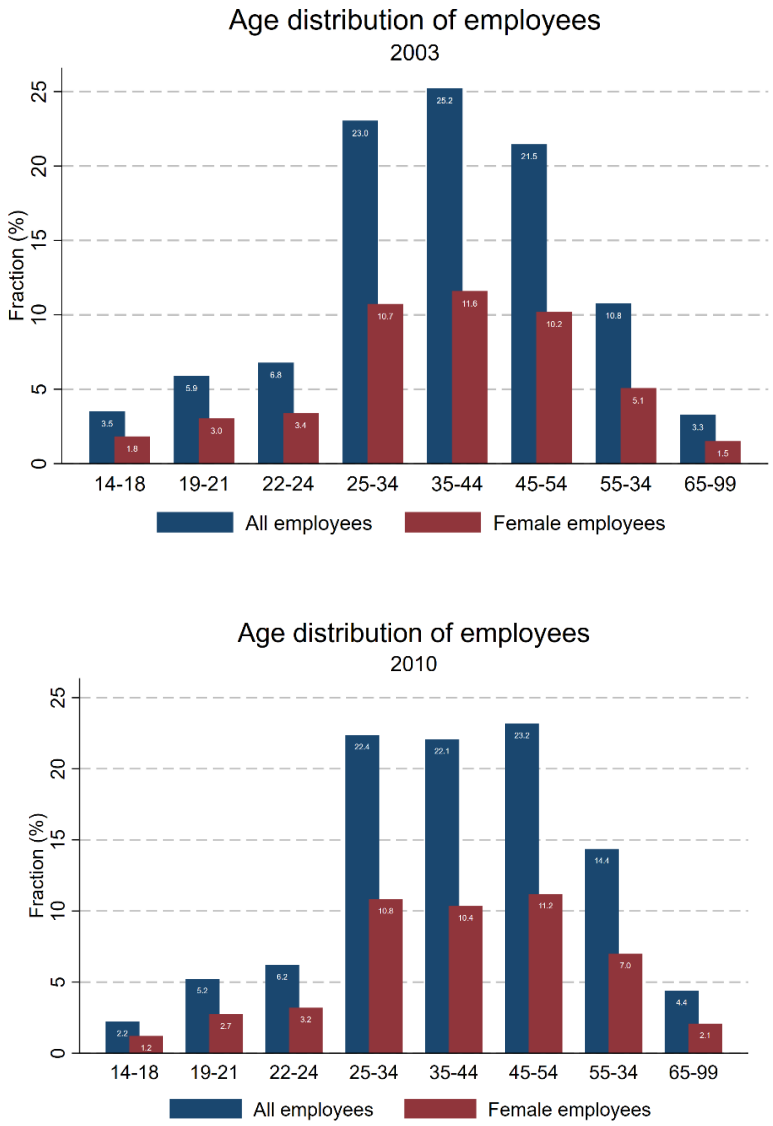
Table A4. Falsification test--Estimates of California paid family leave in other TDI states

Panel A: All firms	Maternity Leave	Family Leave	Other Leave	Any Leave
Comparison group 1: Mothers of youngest children aged 5 to 17	1.90	2.05	0.13	2.03
	(2.29)	(2.46)	(1.07)	(2.44)
Comparison group 2: Women with no children	2.00	3.08	1.52*	3.52
	(2.31)	(2.37)	(0.83)	(2.50)
Comparison group 3: Mothers of youngest children aged <1 year in FL, PA, TX	1.77	1.83	-0.17	1.60
	(2.73)	(2.55)	(0.92)	(2.51)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA, HI, NY, RI	-0.20	0.62	0.77	0.57
	(2.45)	(2.54)	(0.84)	(2.50)
Panel B: Large firms				
Comparison group 1: Mothers of youngest children aged 5 to 17	0.59	2.21	0.67	1.27
	(2.71)	(3.26)	(1.59)	(3.22)
Comparison group 2: Women with no children	1.03	3.63	2.68**	3.71
	(2.84)	(3.14)	(1.14)	(3.14)
Comparison group 3: Mothers of youngest children aged <1 year in FL, PA, TX	0.15	1.86	0.97	1.12
	(2.81)	(2.98)	(1.17)	(3.14)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA, HI, NY, RI	-2.22	0.13	1.88	-0.35
	(2.86)	(3.23)	(1.27)	(3.17)
Panel C: Small firms				
Comparison group 1: Mothers of youngest children aged 5 to 17	5.71*	3.22	-1.58	4.13
	(3.12)	(3.18)	(1.86)	(3.12)
Comparison group 2: Women with no children	5.54	3.57	-0.84	4.69
	(3.11)	(3.20)	(1.92)	(3.18)

Comparison group 3: Mothers of youngest children	5.89	2.57	-2.46	3.43
	(5.02)	(4.33)	(2.23)	(4.27)
Comparison group 4: Mothers of youngest children aged <1 year, all states except CA, HI, NY, RI	3.87 (3.45)	1.80 (3.47)	-1.14 (1.96)	2.73 (3.53)

Note: Following Rossin-Slater, Ruhm, and Waldfogel (2013), Table 5: Each coefficient is from a separate regression, with standard errors in parentheses. The data comes from the 1999 to 2010 March CPS Surveys. All estimates are reported as percentages. The sample is limited to women in the adult civilian population ages 15-64 who reported working any usual hours in the last year. The placebo treatment group consists of women with a youngest child aged <1 who reside in Hawaii, New York, or Rhode Island (these states offer Temporary Disability Insurance, but no access to paid leave during the sample period). The comparison groups include (1) women with a youngest child aged 5 to 17 years in these states; (2) women with no children in these states; (3) women with a youngest child aged <1 year in Florida, Pennsylvania, or Texas (replacing New York with Pennsylvania in this specification as New York is a TDI state); (4) women with a youngest child aged <1 year in all states except California and those in the placebo treatment. All regressions include controls for age buckets (<20, 20 to 29, 30 to 39, 40 to 49, 50 to 59, 60+), indicators for race and ethnicity (non-Hispanic white, black, Hispanic, other), indicators for marital status (married, divorced, separated, widowed, never married), an indicator for being born in the United States, indicators for education categories (< HS, HS, some college, college or more), and indicators for single years of the age of the youngest child. All regressions are weighted by March CPS person weights. State fixed effects are included in these specifications. Maternity leave indicates mothers who were employed and respond that they were absent from the job due to maternity leave last week; family leave includes these mothers plus mothers absent from work due to vacation and personal days, child care obligations, or other reasons. Other leave includes all reasons for job absences except for maternity leave. Any leave includes mothers employed but absent from work for any reason. The large firm subsample includes women who report working at firms with more than 100 employees; the small firm subsample includes women who report working at firms with 0-25 employees.

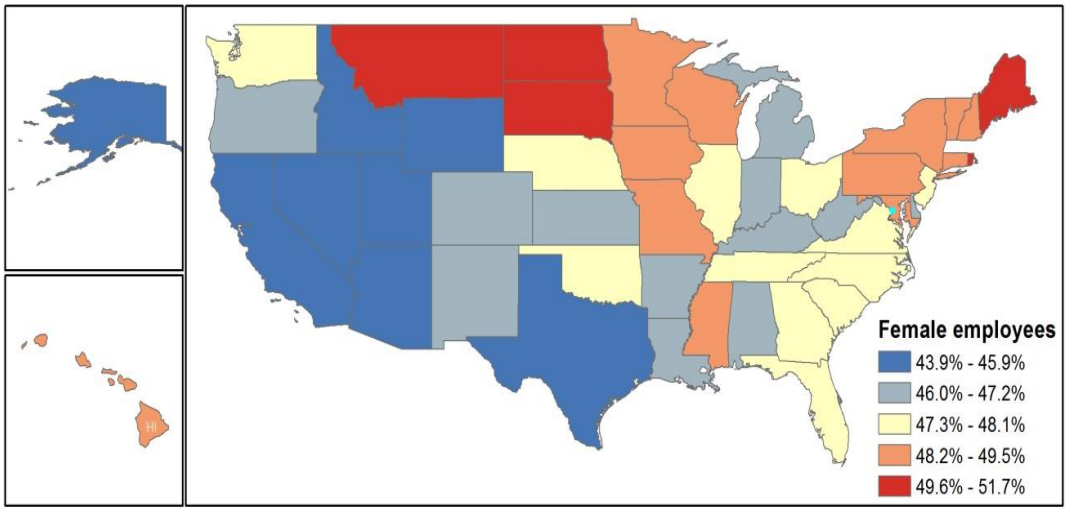
Figure A1. Age Distribution of Employees in 2003 and 2010



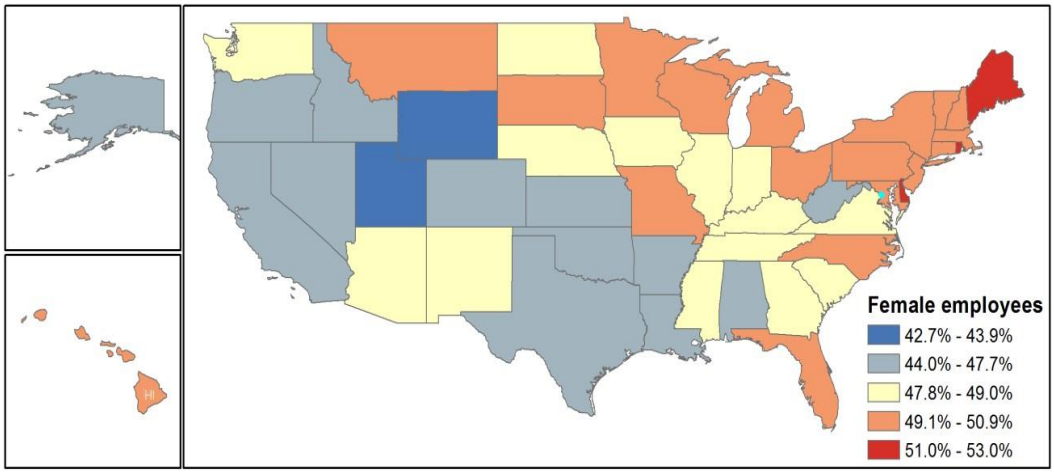
Note: Data are for employed civilians from the Census LEHD for each age group category above. These graphs report average ages for employees at two distinct time periods: 2003, before the enactment of paid leave in California and 2010, after the enactment of paid leave.

Figure A2. Geographic Distribution of Female Employees in 2003 and 2010

2003



2010



Note: Data are for employed civilians from the Census LEHD for each state. These maps report percentage of female employees at two distinct time periods: 2003, before the enactment of paid leave in California and 2010, after the enactment of paid leave.