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The Effects of Paid Family Leave in California on Labor Market Outcomes

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<u>Abstract</u>

Using data from the 1997-cohort of the National Longitudinal Survey of Youth (NLSY-97), we examine the effects of California's first in the nation government-mandated paid family leave program (*CA-PFL*) on mothers' and fathers' use of leave during the period surrounding child birth, and on the timing of mothers' return to work, the probability of eventually returning to pre-childbirth jobs, and subsequent labor market outcomes. Our results show that *CA-PFL* raised leave-taking by around 2.4 weeks for the average mother and just under one week for the average father. The timing of the increased leave use – immediately after birth for men and around the time that temporary disability insurance benefits are exhausted for women – is consistent with causal effects of *CA-PFL*. Rights to paid leave are also associated with higher work and employment probabilities for mothers nine to twelve months after birth, possibly because they increase job continuity among those with relatively weak labor force attachments. We also find positive effects of California's program on hours and weeks of work during their child's second year of life and possibly also on wages.

Key Words: Paid Leave, Family Leave, Employment, Wages, Leave-Taking, Return-to-Work Decisions.

JEL Codes: J1, J2, J3

I. Introduction

Most industrialized countries provide new mothers (and sometimes fathers) rights to a substantial amount of paid leave following the birth of a child. For example, German mothers may take up to a year off from work while receiving 67 percent of their usual pay, and Canada provides a year or more of maternity leave with 55 percent of pay replaced. Conversely, the United States is one of only four nations without entitlements to paid leave (Heymann, Earle, and Hayes, 2007).

Prior to the 1993 Family and Medical Leave Act (FMLA), the U.S. did not provide federal rights to *unpaid* leave either.[1] However, just as some states passed their own laws granting unpaid maternity leave before the FMLA, states have begun to provide paid family leave (*PFL*) from work to care for a newborn or a sick child, spouse, or parent. California was the first state to do so, approving six weeks of *PFL* with 55 percent of usual pay replaced (up to $1,067 per week in 2013), although this leave is not job-protected and is typically not provided to public-sector employees.[2]

California's paid family leave statute (*CA-PFL*), which was passed in 2002 and took effect July 1, 2004, is financed through a payroll tax levied on employees and was added to the pre-

[1] The FMLA provides for 12 weeks of unpaid leave following the birth or adoption of a child, with exemptions for small firms and employees not meeting a work history requirement. The law also covers time off work due to their own or a family member's serious health problem, and so is called "family leave" rather than "parental leave." Along this dimension, the FMLA and the state laws we discuss below are broader than the provisions in many other countries. See Ruhm (2011) for a detailed discussion of family and parental leave laws in both a U.S. and an international context.

[2] Information on California's paid leave program in this and the next paragraph is obtained from Fass (2009); Applebaum and Milkman (2011) and Employment Development Department (2013).

existing Temporary Disability Insurance program that typically provides mothers with six weeks of paid leave during or just after pregnancy. In July 2009, New Jersey began a "family leave insurance" program quite similar to *CA-PFL*, also added to the state's TDI system, which offers six weeks of paid leave at a 66 percent replacement rate, although with a considerably lower ($584 per week in 2013) maximum benefit (Department of Labor and Workforce Development, 2013). Beginning in 2014, Rhode Island's "temporary caregiver's insurance" program will provide four weeks of paid leave at a 60 percent wage replacement rate, up to a ceiling ($752 per week in 2014). As with California and New Jersey, the program is coordinated with the state's temporary disability insurance; however, job protection is also provided during the leave period.[3] Washington state approved $250 per week in paid benefits to be provided for five weeks, with the program scheduled to begin in 2009 (Progressive States Network, 2010); however, due to budgetary pressures, implementation has been repeatedly postponed and is now not scheduled until 2015 (Employment Security Department, 2013).[4] In addition, President Obama proposed (unsuccessfully) in his 2011

[3] Information on the Rhode Island program is available at www.dlt.ri.gov/tdi/tdifaqs.htm and www.shrm.org/LegalIssues/StateandLocalResources/Pages/Rhode-Island-Temporary-Caregiver-Leave.aspx.

[4] Unlike California and New Jersey, Washington does not have a temporary disability system upon which paid family leave could be added. Only three other states – Hawaii, New York, and Rhode Island – have temporary disability insurance programs (Fass, 2009) and the TDI benefits are often quite low (e.g. the maximum benefit in New York is $170 per week in 2013).

budget, to allocate $50 million in competitive grants to states that start *PFL* programs and there have been increasing efforts to establish a national paid leave program.[5]

Researchers have previously analyzed the labor market effects of (largely) unpaid family leave in the United States (Klerman and Leibowitz, 1997, 1999; Waldfogel, 1999; Baum, 2003a,b; Han and Waldfogel, 2003; Berger and Waldfogel, 2004; Han, Ruhm, and Waldfogel, 2009) and of paid parental leave in other industrialized countries (Ruhm and Teague, 1997; Albrecht et al., 1998; Ruhm, 1998; Ondrich et al., 1999; Schonberg and Ludsteck, 2007; Baker and Milligan, 2008; Gupta, Smith, and Verner, 2008; Hanratty and Trzcinski, 2009; Lalive and Zweimuller, 2009; Pronzato, 2009). These studies typically examine the effects of the government mandates on aggregate employment rates or wages of mothers or women of childbearing age.[6] Most of this research suggests that parental leave rights yield positive effects on labor market outcomes, but with some variation in the findings. For example, Ruhm (1998) indicates that short- to medium-length leave

[5] Most recently, the Family and Medical Insurance Leave Act, proposed by Sen. Kirsten Gillibrand and Rep. Rosa DeLauro in 2013, would provide workers with 12 weeks of paid leave at a 66 percent wage replacement rate (up to a ceiling), with no employer size exemption, and administered by a new Office of Paid Family and Medical Leave within the Social Security Administration (www.nationalpartnership.org/research-library/work-family/paid-leave/family-act-fact-sheet.pdf).

[6] Parental leave rights could increase aggregate employment and wage levels because they preserve employer-employee relationships. Conversely, they may have the opposite effect (for at least some groups) if they raises labor costs (particularly for the workers most likely to take leave). There is also a related literature examining how parental leave entitlements affect the mental or physical health of children and parents (e.g. Ruhm, 2000; Chatterji and Markowitz, 2005; Tanaka, 2005; Berger et al., 2005; Baker and Milligan, 2010; Rossin, 2011).

mandates in Europe increase employment without decreasing wages, whereas Lalive and Zweimuller (2009) find that an extension of Austrian paid leave rights from one to two years decreased maternal employment and wages in the short-term but not the long-run.

Paid family leave could have different consequences than the unpaid leave provided under the 1993 FMLA because wage replacement may allow parents facing financial constraints to take more time off work. Moreover, coverage under California PFL is nearly universal, whereas fewer than 60 percent of workers are eligible under the FMLA, due to its firm size and work history requirements.[7] The effects of the California paid leave program may also depart from those of paid leave in other industrialized nations because of its relatively short duration (e.g., six weeks in California versus a year or more in Canada).[8]

PFL is expected to raise leave-taking in the period immediately following the birth because some parents will delay their return to the pre-childbirth job, during which time they are "employed but not at work," while others take leave rather than quitting their jobs. However, to the extent that job continuity is increased, employment and work may rise in the longer-term. These effects will be

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[7] Klerman et al. (2012) estimate that 59 percent of workers were FMLA-eligible in 2012. Eligibility rates will be lower for expectant parents who work for smaller firms or have less recent employment experience than the average worker.

[8] Benefits in some European countries are long enough to allow parents to have multiple births while on paid leave.

dampened to the extent that parents have paid leave even without the legislation or if the wage replacement rate is too low for them to afford time off work.[9]

Most closely related to the current research is Rossin-Slater, et al.'s (2013) analysis of March Current Population Survey (CPS) data from 1999-2010, which shows that *CA-PFL* more than doubled the use of maternity leave among mothers with infants – increasing it from three to six or seven weeks for the average mother. They also provide suggestive evidence of particularly large growth in use for less advantaged groups and of medium-term increases in the usual weekly work hours and wages of employed mothers of one to three year old children. However, the March CPS does not identify the precise timing of leave-taking nor permit testing of whether the increases in leave use occurring during the period in which *CA-PFL* is anticipated to have the strongest effects.

We build on the existing literature by using the data from the 1997 cohort of the National Longitudinal Survey of Youth (NLSY-97) to examine how *CA-PFL* affected leave-taking and (for mothers) other labor market outcomes. The NLSY-97 provides information on the location and exact timing of births, as well as detailed work history data before and after it. Our analysis focuses on parents with substantial work experience during the pregnancy period, since this is the group potentially eligible for paid parental leave, and uses a differences-in-differences (DD) approach where the experiences of new California parents in the period after *CA-PFL* implementation are compared to their counterparts whose children were born earlier, and these changes are contrasted with corresponding parents in matched comparison states. Our analysis extends the literature by examining parents' leave and work decisions in each day and week after the child's birth, by

[9] In 2012, 35 percent of female employees were at worksites offering paid maternity leave (although of potentially short duration) to "all" or "most" employees and 20 percent of males were at sites offering corresponding paternity leave (Klerman, et al., 2012).

investigating the likelihood and timing of the return to the pre-birth job, and by analyzing fathers (to the extent the data allow), as well as mothers.[10]

Since the NLSY-97 precisely identifies the timing of births and leave-taking, we are able to determine whether the patterns of leave use are those anticipated by the institutional details of *CA-PFL*, making a more credible case for causal inference. Specifically, California mothers are expected to begin using PFL following the exhaustion of temporary disability benefits, which typically occurs six to eight weeks after birth, so that this is where we should see increases in leave-taking following enactment of the program. Conversely, since fathers do not qualify for (pregnancy-related) TDI benefits, any increase for them should occur immediately after the birth. Examining subsequent employment rates – a year or more after the child's birth—shows longer-term effects of the government leave mandates (e.g. occurring through changes in employer-employee relationships).

Our analysis yields six primary results. First, the availability of *CA-PFL* increases leave-taking. On average, mothers use two to three additional weeks of leave and fathers just under one extra week. Second, the timing of the rise in leave use – just after the birth for fathers and around the time temporary disability benefits are exhausted for mothers – is consistent with those expected if the program has a causal effect. Third, the increase in maternal leave use primarily reflects work reductions during the first three months after birth, although some specifications also suggest increased rates of employment. Fourth, the California paid leave program is associated with greater probabilities that mothers have returned to work by mothers nine to twelve months after giving birth.

[10] Han et al. (2009) analyzed unpaid leave in the U.S. includes fathers, but they do not know the exact timing of the birth or the reason why parents are employed but not working. Hanratty and Trzcinski (2009) model the return to work after childbirth among Canadian mothers but cannot distinguish between periods of paid and unpaid leave or nonparticipation in the labor force.

Fifth, the results for job continuity are mixed, providing little evidence of an increased likelihood that mothers return to their pre-birth employers in our main specifications, but with such a pattern emerging when broadening the sample to include those working fewer weeks during pregnancy. This raises the possibility that *PFL* provides incentives for some pregnant women to stay on the job until birth in order to qualify for paid leave benefits (and then subsequently return to the same employer). Last, we find evidence that California's paid leave legislation has positive medium-term effects on weeks and hours worked by mothers and, possibly, also on wages.

II. Data

In 1997, the National Longitudinal Survey of Youth began annually collecting information on the labor market experiences and background characteristics of 4,385 females and 4,599 males aged 12 to 16, including oversamples of non-Hispanic blacks and Hispanics. To construct our analysis sample, we selected respondents (mothers and fathers) who had a child between 2000 and 2010 (the last available survey wave) and, in our main specifications, who were employed at least 32 weeks during the nine months before the child's birth. We excluded parents who worked fewer weeks during the pregnancy, since they would be unlikely to qualify for family leave, although we test and report on the robustness of the results to this exclusion.[11] We also omitted the self-employed who, by definition, do not need to negotiate for leave from work. When weighted, our sample is nationally representative of children born to parents meeting these conditions.

The NLSY-97 collects weekly data on labor market status (i.e. employed, unemployed, out-of-the labor force), as well as the exact dates of childbirth. This allowed us to construct a work

[11] Parents quitting jobs prior to giving birth will not receive leave benefits. However, some who worked less than 32 weeks during the pregnancy could receive paid leave – e.g. a mother who worked continuously during the second and third but not the first trimester.

history for each mother and father identifying whether she or he was employed in each week before and after the birth. The starting and stopping day of paid and unpaid leave spells (during which the individual was employed but not working) are also identified. The “paid leave” questions refer to paid time off work because of a pregnancy or birth of a child. The “unpaid leave” questions indicate unpaid vacation or leave that is related to a pregnancy (for spouses in the case of men). One complication is that some respondents may classify time off work provided under *CA-PFL* as “unpaid leave” because they are not directly paid by their employer, but instead by the State of California.[12] Therefore, we focus below on the total leave-taking, including both paid and unpaid time off work.

The NLSY-97 questions are designed to capture leaves lasting at least seven days. However, many mothers and fathers report leaves of six or fewer days, and durations of exactly seven days occur only slightly more often than those lasting six or eight days.[13] For this reason, our analysis proceeds as if all leaves are identified, but we recognize that some short leaves are probably missing. This is likely to be particularly problematic for fathers, who frequently will be off work for only

[12] In regression models that separate them, *CA-PFL* is associated with higher use of both paid and unpaid leave, which is consistent with the hypothesized classification problem, since there is no reason why paid leave would raise unpaid time off work.

[13] For example, 2.0%, 3.1%, 2.2%, 4.1% and 2.6% of 2010 year unpaid leaves for female respondents were reported to last four, five, six, seven and eight days respectively. For paid leaves, the corresponding percentages were 1.3%, 0.0%, 0.9%, 1.7%, 1.3%. For fathers, these reported percentages were 5.7%, 1.3%, 1.8%, 7.9%, and 3.1% for unpaid leaves and 11.4%, 5.3%, 8.3%, 19.7%, and 14.4% for paid leaves.

brief periods of time. And the restriction should be noted when we consider leave survival probabilities, which will be overstated by the exclusion of some short leaves.

Mothers are followed for one year after giving birth, in most of the analysis, but the return to work is treated as an absorbing state, so that the tracking is discontinued once this occurs. We are also able to identify the last job held by the mother before her child's birth and the first job after it, and so can determine if she returned to the pre-birth employer. Our analysis of fathers is limited to leave-taking – we do not examine future employment probabilities because it seems unlikely that these will be much affected by the brief leaves that (some) fathers take and because any such effects are likely to be overwhelmed, in the differences-in-differences (DD) framework, by small disparities in levels or trends between the treatment and control groups.

The key explanatory factor, *CA-PFL*, is a dummy variable equal to one if the child is born on or after July 1, 2004, when California's paid family leave is in force, and zero for births in other states or in California before that date. The *CA-PFL* variable is a good indicator of eligibility because coverage is almost universal for private-sector employees. However, some new parents may be ineligible because they stopped working earlier in the pregnancy or do not meet the (weak) work history requirements for coverage.[14] We also control for the parent's age with a comprehensive set of dummy variables (one for each year of age in the sample), race/ethnicity (black and Hispanic), education (years of school completed), marital status (married vs. unmarried), and years of prior

[14] To be eligible, new parents must have earned at least $300 during the 5 to 17 preceding months; there are no other work history or tenure requirements (Applebaum and Milkman, 2011).

work experience. Additional covariates include family size, number of biological children and parity (child birth order).[15]

In supplemental analyses, we explore longer-term effects of CA-PFL on wages and on annual weeks and weekly hours of work. Specifically, we use the NLSY-97's work history data described above to identify the hourly wage at the mother's job held one year after the birth, as well as the number of hours and weeks worked during the child's second year of life (e.g., the 53rd through 104th weeks after birth).[16]

III. Empirical Specification

We use multivariate differences-in-differences (DD) analysis to explore the effects of *CA-PFL*, distinguishing between the impacts on mothers and fathers through the use of separate models for each. Our DD models contrast changes in the outcomes for new California parents before and after enactment of *PFL* to those for corresponding parents in matched control states.

The basic DD specification takes the form:

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[15] We could not further stratify the unmarried group because only 7.2% of the sample were widowed, separated or divorced. Work experience is calculated by summing weeks worked, excluding weeks on unpaid or paid leave, through the week preceding the child's birth and then dividing by 52.

[16] More precisely, wages are measured at the first job held between the 47th to 57th weeks after birth, adjusted for inflation to year-2012 dollars using the Consumer Price Index (CPI). We also windsorize wages, replacing values below (above) $5 ($50) with $5 ($50). Windsorizing at other values (e.g., $1 and $100) does not appreciably affect the results. The measures of work weeks and hours do not condition on employment – i.e. they include zero values.

$$Y_{ict} = \alpha + \beta_1 CA_{ict} + \beta_2 POST_{ict} + \beta_3 CA{\times}POST_{ict} + \boldsymbol{\gamma_1 X}_{ict} + \varepsilon_{ict}, \qquad (1)$$

where *Y* is the outcome, *CA* is dummy variable taking the value of one for California parents and zero for their control state counterparts, *POST* is a dichotomous indicator set to one (zero) for births on or after (before) the July 1, 2004 enactment of *PFL*, ***X*** is a vector of supplementary covariates, ε is an error term, and the subscripts respectively denote parent *i*, child *c*, and *t* days or weeks after the birth.[17] $\hat{\beta}_3$ provides the DD estimate of primary interest. We obtained (but do not show) similar results using the somewhat more flexible model:

$$Y_{ict} = \alpha + \beta_3 CA{\times}POST_{ict} + \boldsymbol{\gamma_1 X}_{ict} + \boldsymbol{\gamma_2 T}_{ict} + \boldsymbol{\gamma_3 S}_{ict} + \varepsilon_{ict}, \qquad (1')$$

where ***T*** and ***S*** are vectors of year and state dummy variables.[18]

The outcomes are measures of labor market status including the use of leave before or after birth and, for mothers, the probability of work, employment, and having returned to the pre-birth employer. As mentioned, we restrict the main analysis to parents who worked at least 32 weeks during the pregnancy, since those who have not done so are unlikely to qualify for leave.[19] The tables also report robust standard errors, clustered at the state level (Bertrand, Duflo, and Mullainathan, 2004).

We present additional results using several variants of (1). In some figures, we show component elements of the DD models through visual comparisons of the outcome variables for

[17] Negative values for *t* indicate periods before the birth.

[18] (1') controls for a more complete set of time-invariant location-specific effects and for factors that vary uniformly across locations at a point in time. The *POST* and *CA* main effects are absorbed by ***T*** and ***S***, and so do not show up in (1').

[19] This sample inclusion criterion also excludes older women and single men.

California and comparison state parents before and following the implementation of *PFL*, after controlling for demographic characteristics. This is done through estimates of:

$$Y_{ict} = \alpha + \beta_1 CONTROL\times POST_{ict} + \beta_2 CA\times PRE_{ict} + \beta_3 CA\times POST_{ict} + \boldsymbol{\gamma_1 X}_{ict} + \varepsilon_{ict}, \qquad (2)$$

where *CONTROL* is a dummy variable equal to one for births from the control states and zero for births from California, and *PRE* is a dichotomous variable set to one for births before the July 1, 2004 implementation of *PFL* implementation. In (2), $\hat{\alpha}$ provides the regression-controlled estimated average value of the dependent variable for the reference group of control state parents prior to *PFL*. Corresponding estimates for control state parents after July 1, 2004 and California parents before and subsequent to *PFL* implementation are $\hat{\alpha} + \hat{\beta}_1$, $\hat{\alpha} + \hat{\beta}_2$, and $\hat{\alpha} + \hat{\beta}_3$ respectively.

We also sometimes estimate parental leave hazard and survival rates for specified periods after birth. This is done using discrete time hazard models (Prentice and Gloeckler,1978; Meyer, 1990, 1995) measuring the probability that a spell of leave ends between week (or day) *t* and *t+1*, conditional on being on leave at *t*.[20] Hazard models are conceptually appropriate and are well designed to deal with censored observations (e.g., when some leave spells are on-going as of the most recently-released wave of data) and the discrete time specification imposes no parametric restrictions on the underlying baseline hazard function.

Defining $\lambda_{ic}(t)$ as the hazard rate *t* weeks (or days) after the birth of child *c* for parent *i*,

$$\lambda_{ic}(t) = prob[t+1 \geq T_{ic} \mid T_{ic} \geq t], \qquad (3)$$

and the hazard specification is

$$\lambda_{ic}(t) = \lambda_0(t) e^{\beta_1 X_{ict}}, \qquad (4)$$

[20] In our application, once a mother (or father) moves off leave, either by returning to work or exiting the labor force, she exits the sample.

with $\boldsymbol{X}_{ict}$ defined as above, except with an additional control for an interaction of *PFL* with a quartic function of leave duration. $\lambda_0(t)$, the baseline hazard rate for week (or day) t, is estimated non-parametrically.[21] The corresponding survivor rate, $\Phi_{ic}(T)$, which is the probability of remaining on leave for T weeks (or days), is the cumulative product of (one minus) the individual hazard rates, or

$$\Phi_{ic}(T) = \prod_{t=1}^{T}(1-\lambda_{ic}(t)). \quad (5)$$

IV. Selection of Control States

A requirement for the DD procedures to generate consistent estimates of the causal effect of *CA-PFL* is that the changes over time in the outcomes would have been similar between California and the control states had *CA-PFL* not been enacted, although the levels could differ. Conversely, if, for example, leave-taking was increasing faster for California mothers than for counterparts in other states, even absent *CA-PFL*, then the program will be spuriously related to increases in leave use, leading to an overestimate of its true causal effect.

While we cannot know what the outcomes during the post-*PFL* period in California would have been without its implementation, we can observe the pre-program trends. Therefore, our empirical strategy is to choose control states with similar trends in maternal leave-taking before July 1, 2004 to those observed in California. Specifically, we use the following procedure to determine whether parents in a given state should be included in the control group. First, we exclude the 29 states and the District of Columbia with fewer than 8 NLSY-97 women giving birth during the pre-program period, since precision of the estimates will be extremely low in these cases. Second, for

[21] Specifically, we use the probit functional form (see Maddala, 1983) for $\lambda_{ic}(t)$, including duration dummy variables which allow the baseline hazard to take a value in each period that best fits the data, instead of being forced to follow a trend that is partially determined by other durations.

each of the remaining 21 states, we estimate the following model, using only observations from before July 1, 2004 for California and the specified state:

$$Y_{ict} = \alpha_0 + \alpha_1 \boldsymbol{X}_{ict} + \alpha_2 TR_{ict} + \alpha_3 NONCA_{ict} + \alpha_4 TR \times NONCA_{ict} + \varepsilon_{ict}. \qquad (6)$$

In (6), *TR* is a linear time trend (for the 2000-2004 period) and *NONCA* is a dummy variable set to one for the potential control state and zero for California. The interaction term, $TR \times NONCA_{ict}$, allows the leave-taking time trend to differ between California and the other state, and we treat that state as a possible control if we are unable to reject the null hypothesis that $\hat{\alpha}_4$ equals zero. Although similar numbers of NLSY-97 fathers and mothers have children before July 2004, too few fathers take leave to identify valid control states. Therefore, we use the control states identified for mothers in our analysis of the effects of *PFL* on father's leave-taking.[22]

Table 1 presents $\hat{\alpha}_4$ from equation (6) for each potential control state, the associated standard error, the number of pre-July 2004 births to mothers, and whether each state is included in the control group. Ultimately, 15 states are deemed to be valid controls for mothers. The other 35 states and District of Columbia either had different pre-*PFL* trends in leave-taking or provided too few observations to be compared with California. In addition to eliminating those states with statistically different time trend coefficients, we exclude six states (AL, DE, MS, MO, TN, and WI) whose time trend coefficients are not statistically different but exceed 1.0 in absolute value. Although admittedly arbitrary, we consider the pre-trend patterns in these states to be too different from those in California to provide valid controls.

[22] We obtain similar marginal effects of *CA-PFL* on leave-taking, with somewhat smaller standard errors in some cases, for fathers when using all non-California states as controls.

Descriptive characteristics, weighted so as to be representative, are provided for mothers and fathers in Appendix tables A.1 and A.2, with separate results presented for California and the control states and for periods before after *CA-PFL* implementation. The combined (California plus control state) samples contain 1,188 births for mothers and 1,126 births for fathers. California parents are less likely to be black and more often Hispanic than parents from control states. As expected, since the NLSY-97 follows a cohort, those giving birth before July 2004 are younger, have less education and work experience, are less likely to be married, and have fewer children than counterparts whose children are born later. However, these parents reside in households with more members.

V. Leave-Taking

Figure 1 presents the daily regression-adjusted proportion of mothers in our sample on leave during the 12 weeks before and 39 weeks after giving birth. These are obtained from estimates of equation (2) which, as discussed, distinguish births before and after July of 2004 and in California versus the control states, and adjusts for differences in race/ethnicity, age, education, marital status, work experience, family size and number of children.

The figure provides strong evidence that *CA-PFL* increased leave-taking. Prior to the program's enactment, new mothers in California took roughly the same amount of leave as their control state counterparts in the weeks immediately before birth and slightly more two to eight weeks after it (probably reflecting the availability of temporary disability insurance in California). There was no change, or possibly even a slight decrease, in control state leave-taking after July of 2004, whereas its use increased fairly dramatically for California mothers over the same period.

Additional details are provided in the first two columns of Table 2, which shows differences-in-differences estimates (the estimated marginal effects on the *CA×POST* interaction in equation 1) along with the associated standard errors, for specified time periods after the births. For example, the first table entry indicates that California's paid leave program raised estimated leave-taking one day after birth by a highly significant 17.0 percentage points (from a baseline of 57.7%). The DD estimates indicate that *PFL* was associated with 14 to 17 percentage point in leave-taking during the first five or six weeks after birth and 18 to 30 point growth during the next seven weeks. The effect shrinks rapidly thereafter, although generally remaining statistically significant through the child's first four months. These patterns make sense since *PFL* is expected to have the strongest effect during the six-week period after the expiration of Temporary Disability Insurance benefits, which generally exhaust for mothers six to eight weeks after the birth.

PFL also appears to have increased the leave-taking of fathers, but with three important differences. First, the strongest effects occur immediately after the birth, rather than being delayed by several weeks (see Figure 2 and the last two columns of Table 2). This is again consistent with a causal effect of *CA-PFL*, since many women will be on TDI leave after delivery, which fathers are not eligible for. Second, the absolute magnitude of the effect is much smaller – peaking at 6 to 10 percentage points for men versus 20 to 30 points for women. However, since the baseline rates of leave-taking are also dramatically lower for fathers (17.0% percent just after birth versus 57.7% for mothers) the estimated effects are of approximately equal size or even slightly larger in relative terms.[23] Third, fathers remain on leave relatively briefly, even after the enactment of paid family leave, with less than 6 percent still off the job by the end of the child's third week and fewer than 1

[23] Applebaum and Milkman (2011) indicate the proportion of fathers taking paid leave after a child's birth in California increased from 17 percent in 2004 to 26 percent in 2010.

percent after the seventh week.[24] By contrast, after the enactment of *PFL*, 53.3% of California mothers and 24.7% of those in control states were still not at work seven weeks after delivery.

Further detail on how *CA-PFL* affected the timing of leave-taking is provided in Table 3 and Figures 3 through 6, which show estimated parental leave hazard and survival rates after controlling for demographic characteristics. Specification 1, in Table 3, constrains the effects to be proportional in each week after birth; in specification 2 and the figures, the effects of *CA-PFL* on the hazard rates are allowed to vary nonlinearly with leave durations: this is done by interacting a quartic polynomial of time on leave with the indicator for the post-*PFL* implementation period (with California and *Post* main effects also controlled for).

CA-PFL reduces the estimated average weekly hazard rate out of leave by a statistically significant 4 to 5 percent for both men and women (see specification 1). However, as shown in specification 2, the predicted effects vary substantially over time. For mothers, the hazard rates fall immediately after the birth but with the largest decrease occurring 6 to 14 weeks after delivery and with a negative effect persisting until around the 18th week, after which the pattern reverses and *CA-PFL* is associated with higher hazard rates (Figure 3). This last effect occurs because fewer than five percent of California mothers remained on leave through the 18th week prior to 2004, versus around 10 percent after the law's passage (see Figure 4), so that hazard rates became necessarily low. Almost no mothers remain on leave beyond six months, either before or after the enactment of California paid leave. For fathers, the reduction in leave hazard rates is immediate, peaking one week after the birth, and with a higher exit rate predicted during the post-*PFL* period after the second

[24] On a related point, we do not show regression-adjusted effects for fathers beyond two weeks, in Table 2, because no California fathers remained on leave after 14 days during the pre-*PFL* period.

post-birth week (Figure 5), reflecting the extremely low rates at which fathers took more than two weeks of leave before 2004 (see Figure 6). These results suggest a causal impact of *CA-PFL*. Specifically, the reduction in expected hazard rates is largest for women during the several weeks after the exhaustion of temporary disability insurance benefits, six to eight weeks after birth, but immediately after it for men, who do not have access to TDI.

Using the predicted survivor rates, displayed in Figures 4 and 6, we estimate that *CA-PFL* raised average leave-taking from 7.8 to 10.2 weeks for new mothers and from 11.0 to 15.7 days for new fathers.[25] This implies that the average mother is taking 40 percent of the statutory duration of the program as additional leave (2.4 of 6 weeks) and that the average father is taking around one-sixth of the newly available leave (4.7 days out of the 6 weeks). In addition to these net increases, there may be some replacement of time off work that would have otherwise been taken as unpaid or company paid leave. These average effects conceal substantial variation in leave use across parents. For instance we estimate that leave increased by around two (four) weeks for mothers at the 25th (75th) percentiles of leave use and by 3 (10) days for corresponding fathers.

VI. Other Labor Market Outcomes

The increase in leave-taking due to *CA-PFL* reflects a small reduction in non-employment combined with a larger decrease in work among those who remain employed (but on leave). This is shown in Table 4, which provides separate estimates for nonemployment and work, as well as for work at the job held prior to childbirth. For instance, at the end of the first post-birth week, the 16.6 percentage point rise in predicted leave-taking (shown in Table 2 and previously discussed), consisted of a 3.9 percentage point decline in non-employment and a 12.6 point reduction in work.

[25] This is calculated as $\boldsymbol{\Sigma}_t \Phi(t) \times t$, for $\Phi(t)$ the probability of being on leave t periods after the birth.

As mentioned, large predicted effects on maternal leave-taking persist through the fourth month after delivery and, at almost all of these intervals, are accompanied by strong reductions in work and much weaker (and usually statistically insignificant) declines in non-employment. Additional detail on the components of the DD estimates – the regression-adjusted labor market status of California and control state mothers before and after *CA-PFL* enactment – are provided in Figures 7 and 8.[26]

The intermediate-term labor market effects are equally interesting. *CA-PFL* is associated with increased leave-taking and reduced rates of work during the first four months or so after birth, as mentioned; however, by month six the leave-taking is complete and the negative predicted effects on work have been eliminated. By nine months after birth, *CA-PFL* is predicted to *increase* work probabilities and to reduce non-employment by a statistically significant five to six percentage points, an effect which persists through at least the end of the first year.

Researchers previously examining (mostly unpaid) state and federal leave entitlements (Washbrook, et al., 2011), as well as California's paid family leave program (Rossin-Slater, et al., 2013) have also found that leave rights initially reduce but subsequently increase rates of work. The reason generally hypothesized for the positive intermediate-term effect is that the availability of leave reduces quits and raises the probability that mothers remain with their pre-birth employer. We explore this possibility in the last two columns of Table 4 and in Figure 9, where the dichotomous outcome indicates whether the mother is working at the last job held prior to giving birth.

[26] Figure 7 also shows the origin of the increase in non-employment estimated to occur in week 8. Notice that the predicted non-employment rates of pre-*PFL* control group mothers spike downwards and those of corresponding California mothers spike upwards, in this one week, before reverting to more normal levels in the next. This produces the week 8 estimate shown on Table 4.

Our main specifications provide little evidence that *CA-PFL* increased job continuity. Specifically, California mothers are predicted to be around one percentage point more likely to work at their pre-birth job nine or twelve months after delivery, after enactment of the paid leave program, but the effect does not approach statistical significance. The reason the impact is so small is that the vast majority of the mothers analyzed (over 80 percent for all groups) eventually return to their old jobs. However, this could partially reflect our sample inclusion requirement of having worked at least 32 weeks during the pregnancy. Specifically, the availability of paid leave might induce some mothers to work more during this period and take short paid leaves after it, rather than quitting the job held during pregnancy, thereby increasing job continuity. This possibility is examined below.

VII. Robustness Checks

We next test whether the preceding results are robust to changes in the choice of control states or sample inclusion criteria. Table 5 summarizes the results. The top panel shows our base model, where the comparison group consists of the matched control states and the sample includes mothers employed at least 32 weeks of the pregnancy period. The second panel expands the comparison to include all states (not just the matched controls). The third and fourth panels return to the control state comparison but reduce the pre-birth work requirement for inclusion in the analysis to 20 weeks and any employment during the nine months before the birth.

Most results are insensitive to these changes. Expanding the control group to include all states has little impact. Weakening the pregnancy period work requirement does not materially or consistently affect the results for leave-taking but changes the estimated *CA-PFL* effect on nonemployment and work in two ways. First, higher levels of employment explain a greater portion of the increase in leave-taking, with a consequent decrease in the contribution of reductions in work. For instance, *CA-PFL* is predicted to reduce the nonemployment of mothers by 7.8 percentage points, 4 weeks after the birth, when the sample includes those working at least 20 weeks during pregnancy, versus a 1.3 point decrease when restricting the analysis to those who worked at least 32 weeks during this period. The accompanying reduction in work probabilities is 12.0 rather than 15.7 percentage points. Second, the medium-term increases in employment, work and return to the pre-birth job are considerably larger when using less restrictive sampling criteria. Thus, *PFL* is predicted to raise the probability of having returned to work within one year by 10.5 (8.7) percentage points and to have done so with the pre-birth employer by 5.9 (8.8) points among mothers with any (at least 20) weeks of work during pregnancy, compared to 6.2 and 1.1 percentage point increases for the main sample.

This suggests that *CA-PFL* increases the job continuity of new mothers in ways that our main estimates do not capture. Specifically, by restricting the sample to persons with substantial pre-birth employment prior, we may be ignoring reductions in quit rates that paid leave facilitates. Without paid time off the job, some mothers planning to stop working once their children are born may leave their positions before delivery and therefore be excluded from our analysis. However, when paid leave is available, some of them may choose not to quit their jobs but rather to take some time off work and then return to their original employer. This could also help to explain the *PFL*-related reduction in nonemployment observed immediately after childbirth. While an argument could be made for using a less stringent pregnancy work requirement as our main specification, the tradeoff is that doing so is likely to include more parents who who would not continue to be employed to the point where leave rights become a relevant consideration.

VIII. Wages, Earnings, and Work

Last, we examine how California's paid family-leave program has influenced longer-term labor market outcomes of mothers including: the probability of having returned to work within one year of the birth, log hourly wages at this time, and the average number of weeks and hours worked during the second year of the child's life. The analysis samples are smaller than those above for two reasons. First, labor market status will be unavailable for mothers giving birth near the end of the analysis period. Second, wage data are missing for women who are not employed at the end of the relevant time period and are not provided for some working mothers.[27]

Our main results, summarized in the top panel of Table 6, confirm that *CA-PFL* increases rates of maternal work one year after the child's birth. Specifically, the DD estimate suggests that the paid leave program raised the work probabilities of mothers by 5.5 percentage points one year after birth, compared to a pre-*PFL* baseline of 91 percent.[28] This is smaller than the 6.2 point increase obtained in Table 5 (with a slightly larger sample and longer time period) but the difference does not approach statistical significance. Rights to paid leave are also predicted to elevate weeks worked and average weekly work hours during the second year of the child's life – by 6.9 weeks and 4.2 hours – which represent 19 and 16 percent growth compared over the pre-program baselines of 36.8 weeks and 25.9 hours per week. At least some of this increase is expected, since rights to paid leave significantly increased the probability of having returned to work by the end of the child's first year. Finally, the point estimates suggest that hourly wages increase by around five percent, one year after the birth, but the confidence intervals are wide and include zero or negative effects.

[27] We found no evidence of differential patterns in missing data for California versus control state mothers.

[28] These are marginal effects estimated from probit models.

The bottom panel of Table 6 shows that we obtain fairly similar results when broadening the sample to include mothers with any pre-birth employment.[29] The one change is that there is a considerably larger predicted increase in the probability of returning to work by 52 weeks after the birth – 9.6 versus 5.5 percentage points – which is consistent with the results previously described in Table 5 and again suggests a possible *CA-PFL* effect on job continuity.[30]

IX. **Discussion**

Our analysis indicates that California's paid family leave program raised the leave-taking of new mothers and fathers. These increases are sizable and last for four months after the birth for mothers and two weeks subsequent to it for fathers. For mothers, the effects are most pronounced during the sixth through thirteenth weeks after delivery – an estimated 19 to 30 percentage point increase – which corresponds to the period after the exhaustion of temporary disability benefits. Fathers are not eligible for pregnancy-related TDI and the *PFL* effects for them begin immediately after birth and are largest during the following week. Although the overall increase in leave-taking is much smaller than for mothers – between 3 and 11 percentage points during the first two weeks – the baseline levels are also dramatically lower, so that these effects are large in relative terms. We

[29] As expected, the baseline rates of post-birth work are lower for this group, since it includes mothers with weaker attachments to the labor force.

[30] The effect on log wages at the end of year 1 and weeks or hours worked during year 2 are also similar for the other two samples shown in Table 5 (mothers from all states, and those in California and control states working 20 or more weeks during pregnancy), and the patterns of return to work by the end of the first year also correspond with those found there.

estimate *CA-PFL* increased the average leave-taking of mothers by around 2.4 weeks and that of fathers by just under one week.

We also examined other labor market consequences of the paid leave program for mothers. The increased leave-taking immediately after birth results from a combination of reductions in work among the employed and higher rates of employment, with the former being more important in most specifications. There is consistent evidence that *CA-PFL* increased the likelihood that mothers have returned to work by a year after birth and raised maternal hours and weeks of work by 15 to 20 percent during the second year of the child's life. It is also predicted to raise hourly wages at then end of the first year by 5 percent, but this estimate is imprecise and statistically insignificant. Finally, there is some indication that the medium-term increases in the probability of working may reflect increases in job continuity, resulting because paid leave reduces the probability that some expectant mothers quit their jobs prior to giving birth.

It is useful to compare these results to the recent study of California paid leave conducted by Rossin-Slater et al. (2013). In their main specifications, the program is predicted to raise the average leave-taking of eligible mothers by 3.1 to 3.3 weeks, with modestly smaller estimates obtained in some alternative specifications. The similarity of these results to ours is noteworthy given that they use a different data set – the 1999 to 2010 years of the March *Current Population Survey* – that contains a larger sample but less precise information on leave-taking.[31] They also find positive medium-term effects of *CA-PFL* on maternal employment that appear to be of comparable size or somewhat smaller than those that we observe.[32] For example, rights to paid leave are anticipated to raise the weekly work hours of mothers with one-year-old children by 2 to 3 hours per week, versus

[31] Information on leave-taking is only available for the week prior to the survey.

[32] They examine hours worked in the last week and last year for mothers of one to three year olds.

a four hour per week reduction obtained here; however, these differences are statistically indistinguishable. They also (imprecisely) estimate annual earnings increases of around 13 percent for these mothers, largely reflecting growth in predicted work hours rather than hourly wages. Our point estimate is larger – a 16 percent increase in work hours and a 4 to 5 percent rise in hourly wages implies around a 22 percent rise in annual earnings – but again with standard errors that are easily big enough enough to encompass the Rossin-Slater et al. estimate.[33]

Our analysis extends beyond that of Rossin-Slater, et al. (2013) in at least two important ways. First, the detailed NLSY-97 work history data permit us to identify increases in paternal leave-taking that are modest in absolute size (less a week) but large in relative terms, more than a 40 percent increase from a low baseline. Interestingly, such results are consistent with Han, et al.'s (2009) evidence that (largely unpaid) federal and state leave entitlements are associated with 50 percent or larger increases in leave-taking by fathers during the birth month, from extremely low baseline rates.

Second, we are able to more precisely measure the timing of leave-taking. As mentioned, the patterns of leave use – immediately after births for fathers and reaching a maximum shortly after temporary disability benefits are likely to be exhausted for most mothers – suggests that we are observing a causal effect of *CA-PFL*. In addition, using predicted changes in survivor probabilities at different durations, we can estimate the distribution of the rise in leave-taking. Doing so, we find that the 25th percentile of the leave-taking distribution increases from just over 5 weeks to 7 weeks, the 50th percentile from 8 to 11 weeks, and the 75th percentile from about 10 to 14 weeks. Similarly,

[33] The confidence intervals on Rossin-Slater et al. (2013) predictions are also wide and they estimate a larger (21 percent) annual earnings increase for mothers with three year old children.

for fathers the 25th, 50th, and 75th percentiles of the leave-taking distribution increase from 3 to 6, 7 to 12, and 12 to 22 days respectively.

Some analysts advocate expanding California's paid family leave program by further publicizing its existence, raising the wage replacement rate, extending coverage to public sector employees, and providing job protection during the leave.[34] Our results indicate that the program, as currently structured, has appreciable effects on leave-taking and work, so that these extensions are likely to have noticeable effects as well. The cost of *PFL* mandates depends on the frequency and duration with which parents take paid leave, the amount of wages replaced, and administrative costs (including employer and employee accommodations for absent employees). The findings of this analysis can be used to address the frequency and duration components of this cost. Also, the models examining whether mothers return to the pre-birth employers raise the possibility of cost offsets, if *CA-PFL* helps to preserve employer-employee matches and firm-specific human capital. This information is also relevant, for California policymakers set the employee-paid payroll tax rate used to finance the program. Finally, advocates of national paid family leave programs (e.g. O'Leary et al., 2012; Zigler et al., 2012) draw heavily on the California experience and propose to incorporate many of its key features in their proposals, making these findings particularly salient.

[34] In 2009-2010, a majority of California workers in did not know about the *PFL* program and a third of those who knew of it did not apply for the benefits because the wage replacement was too low; others did not take paid leave because they thought they were ineligible or feared that doing so would limit their future potential advancement or result in employment termination (Applebaum and Milkman, 2011).

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Table 1: Results from Regressions to Select Control States

STATE	Number of Pre-July 2004 Births to Mothers	Time-Trend Coefficient	Standard Error	T-Statistic	Included as Control State
AL	8	-1.284	1.192	1.08	No
AZ	14	-0.531	0.532	1.00	Yes
CA	56	N/A	N/A	N/A	Yes
DE	9	-1.221	0.937	1.30	No
FL	13	-0.828	0.853	0.97	Yes
GA	10	0.206	0.676	0.30	Yes
IL	20	-0.550	0.445	1.24	Yes
IN	12	-0.175	0.550	0.32	Yes
MD	10	-0.657	0.670	0.98	Yes
MI	15	0.401	0.671	060	Yes
MN	16	-0.733	0.521	1.41	Yes
MS	10	-1.276	1.083	1.18	No
MO	11	-1.980	1.029	1.92	No
NY	31	0.380	0.454	0.84	Yes
NC	32	0.600	0.459	1.31	Yes
OH	13	-0.046	0.775	0.06	Yes
PA	14	-0.199	0.631	0.32	Yes
TN	13	-1.069	1.110	0.96	No
TX	64	-0.513	0.359	1.43	Yes
VA	28	-0.243	0.468	0.52	Yes
WI	8	1.369	1.132	1.21	No

The time-trend interaction measures the difference in the 2000 to pre-July 2004 time trend between California and each state listed. States with fewer than 8 pre-July 2004 births to mothers are not considered as control states These include (with the number of births shown in parentheses): AK (1); AR (4); CO (4); CT (2); DC (1); HI (0); ID (0); IA (0); KS (4); KY (6); LA (7); ME (0); MA (7); MT (3); NE (1); NV (3); NH (0); NJ (5); NM (2); ND (5); OK (5); OR (4); RI (1); SC (6); SD (3); UT (0); VT (5); WA (3); WV (0); and WY (0). The text provides further criteria for whether other states are included in the control group.

Table 2: Regression-Adjusted Estimated Effects of *CA-PFL* on Leave-Taking

Time Period	Mothers		Fathers	
Day 1	0.170***	(0.050)	0.063**	(0.029)
Day 2	0.172***	(0.049)	0.094***	(0.035)
Day 4	0.165***	(0.048)	0.085**	(0.036)
Day 6	0.158***	(0.048)	0.114***	(0.038)
Day 7	0.166***	(0.048)	0.100***	(0.036)
Day 8	0.166***	(0.048)	0.071**	(0.032)
Day 10	0.169***	(0.049)	0.069*	(0.037)
Day 12	0.174***	(0.046)	0.030	(0.023)
Week 2	0.163***	(0.048)	0.053*	(0.028)
Week 3	0.171***	(0.052)	-	-
Week 4	0.135**	(0.053)	-	-
Week 5	0.161***	(0.052)	-	-
Week 6	0.189***	(0.048)	-	-
Week 7	0.182***	(0.053)	-	-
Week 8	0.295***	(0.058)	-	-
Week 10	0.297***	(0.046)	-	-
Week 13	0.235***	(0.053)	-	-
Week 16	0.114**	(0.046)	-	-
Week 18	0.062*	(0.034)	-	-
Week 20	0.044	(0.027)	-	-
Week 26	0.004	(0.008)	-	-
Week 34	-0.005	(0.003)	-	-
Week 39	-0.004	(0.003)	-	-

Note: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Table shows difference-in-difference estimates (coefficients on $CA{\times}POST$ interactions). The models also control for California and post-July 2004 main effects, as well as race/ethnicity, age, education, marital status, work experience, family size, and the number of biological children. Robust standard errors, clustered at the state level, are in parentheses. There are 1,188 birth observations for mothers and 1,126 birth observations for fathers who were employed in at least 32 pregnancy weeks from California and the control states. No California fathers are on leave beyond 14 days in the pre-PFL period, so estimates are not provided at those durations.

Table 3: Estimated Effects of *CA-PFL* on Hazard Rates Out of Leave

Time Period	Mothers		Fathers	
Specification 1				
PFL	-0.047***	(0.010)	-0.043***	(0.012)
Specification 2				
PFL*Weeks/100	-0.014***	(0.003)	-0.020***	(0.004)
PFL*Weeks^2/1000	0.055	(0.046)	0.215***	(0.058)
PFL*Weeks^3/10000	0.023	(0.018)	-0.066***	(0.022)
PFL*Weeks^4/100000	-0.006***	(0.002)	0.006***	(0.002)

Note: *** p<0.01, ** p<0.05, * p<0.1. Table shows difference-in-difference estimates (coefficients on *CA×POST* interactions), with the same controls as in Table 2. Robust standard errors, clustered at the state level, are in parentheses. Discrete time hazard models are estimated. There are 6,829 birth-week observations from 749 births for mothers and 2,295 birth-day observations from 186 births for fathers in the paid leave hazard model from California and the control states.

Table 4: Regression-Adjusted Estimated Effects of *CA-PFL* on the Labor Market Status of Mothers

Time Period	Not Employed		Returned to Work By		Returned to Old Job By	
Week 1	-0.039	(0.031)	-0.126***	(0.028)	-0.126***	(0.028)
Week 2	-0.038	(0.033)	-0.124***	(0.027)	-0.125***	(0.027)
Week 3	-0.013	(0.032)	-0.159***	(0.027)	-0.157***	(0.027)
Week 4	-0.013	(0.037)	-0.157***	(0.027)	-0.156***	(0.027)
Week 5	-0.007	(0.035)	-0.160***	(0.028)	-0.157***	(0.027)
Week 6	-0.006	(0.033)	-0.190***	(0.036)	-0.190***	(0.036)
Week 7	-0.013	(0.032)	-0.202***	(0.038)	-0.204***	(0.038)
Week 8	-0.077***	(0.021)	-0.198***	(0.049)	-0.206***	(0.048)
Week 10	0.023	(0.038)	-0.308***	(0.047)	-0.325***	(0.044)
Week 13	0.005	(0.029)	-0.203***	(0.036)	-0.229***	(0.036)
Week 16	-0.019	(0.026)	-0.075**	(0.034)	-0.115***	(0.031)
Week 18	-0.015	(0.025)	-0.060*	(0.034)	-0.097**	(0.041)
Week 20	-0.022	(0.026)	-0.031	(0.033)	-0.074*	(0.040)
Week 26	-0.014	(0.021)	0.009	(0.022)	-0.010	(0.032)
Week 34	-0.017	(0.014)	0.023	(0.015)	0.003	(0.026)
Week 39	-0.050***	(0.011)	0.056***	(0.011)	0.012	(0.025)
Week 52	-0.058***	(0.011)	0.062***	(0.010)	0.011	(0.022)

Note: *** p<0.01, ** p<0.05, * p<0.1. Table shows difference-in-difference estimates (coefficients on *CA*×*POST* interactions), with the same controls as in Table 2. Robust standard errors, clustered at the state level, are in parentheses. There are 1,188 birth observations for mothers who were employed in at least 32 pregnancy weeks from California and the control states.

Table 5: Regression-Adjusted Estimated Effects of *CA-PFL* on the Labor Market Status of Mothers, Alternative Samples

Time Period	On Leave		Not Employed		Returned to Work By		Returned to Old Job By	
	Mothers Employed at least 32 Pregnancy Weeks from California and Control States (N=1,188)							
Week 1	0.166***	(0.048)	-0.039	(0.031)	-0.126***	(0.028)	-0.126***	(0.028)
Week 4	0.135***	(0.053)	-0.013	(0.037)	-0.157***	(0.027)	-0.156***	(0.027)
Week 13	0.235***	(0.053)	0.005	(0.029)	-0.203***	(0.036)	-0.229***	(0.036)
Week 26	0.004	(0.008)	-0.014	(0.021)	0.009	(0.022)	-0.010	(0.032)
Week 52	-	-	-0.058***	(0.011)	0.062***	(0.010)	0.011	(0.022)
	Mothers Employed at least 32 Pregnancy Weeks from All States (N=1,763)							
Week 1	0.178***	(0.033)	-0.032	(0.023)	-0.138***	(0.020)	-0.138***	(0.020)
Week 4	0.178***	(0.038)	0.008	(0.027)	-0.182***	(0.022)	-0.181***	(0.022)
Week 13	0.213***	(0.039)	-0.002	(0.022)	-0.187***	(0.031)	-0.226***	(0.030)
Week 26	0.001	(0.005)	-0.012	(0.015)	-0.009	(0.015)	-0.013	(0.023)
Week 52	-	-	-0.052***	(0.008)	0.054***	(0.007)	-0.008	(0.018)
	Mothers Employed at least 20 Pregnancy Weeks from California and Control States (N=1,446)							
Week 1	0.201***	(0.041)	-0.124***	(0.029)	-0.059**	(0.027)	-0.061**	(0.025)
Week 4	0.205***	(0.044)	-0.078**	(0.036)	-0.117***	(0.025)	-0.102***	(0.024)
Week 13	0.218***	(0.052)	-0.047	(0.030)	-0.120***	(0.037)	-0.083***	(0.031)
Week 26	0.005	(0.008)	-0.060***	(0.022)	0.057**	(0.022)	0.073**	(0.029)
Week 52	-	-	-0.083***	(0.013)	0.087***	(0.012)	0.088***	(0.022)
	Mothers Employed at All during the Pregnancy from California and Control States (N=1,893)							
Week 1	0.160***	(0.035)	-0.079**	(0.033)	-0.070***	(0.024)	-0.072***	(0.023)
Week 4	0.158**	(0.036)	-0.051	(0.036)	-0.107***	(0.024)	-0.100***	(0.022)
Week 13	0.173***	(0.043)	-0.043	(0.028)	-0.099***	(0.030)	-0.091***	(0.026)
Week 26	0.005	(0.007)	-0.088***	(0.019)	0.085***	(0.019)	0.046*	(0.028)
Week 52	-	-	-0.102***	(0.013)	0.105***	(0.014)	0.059**	(0.023)

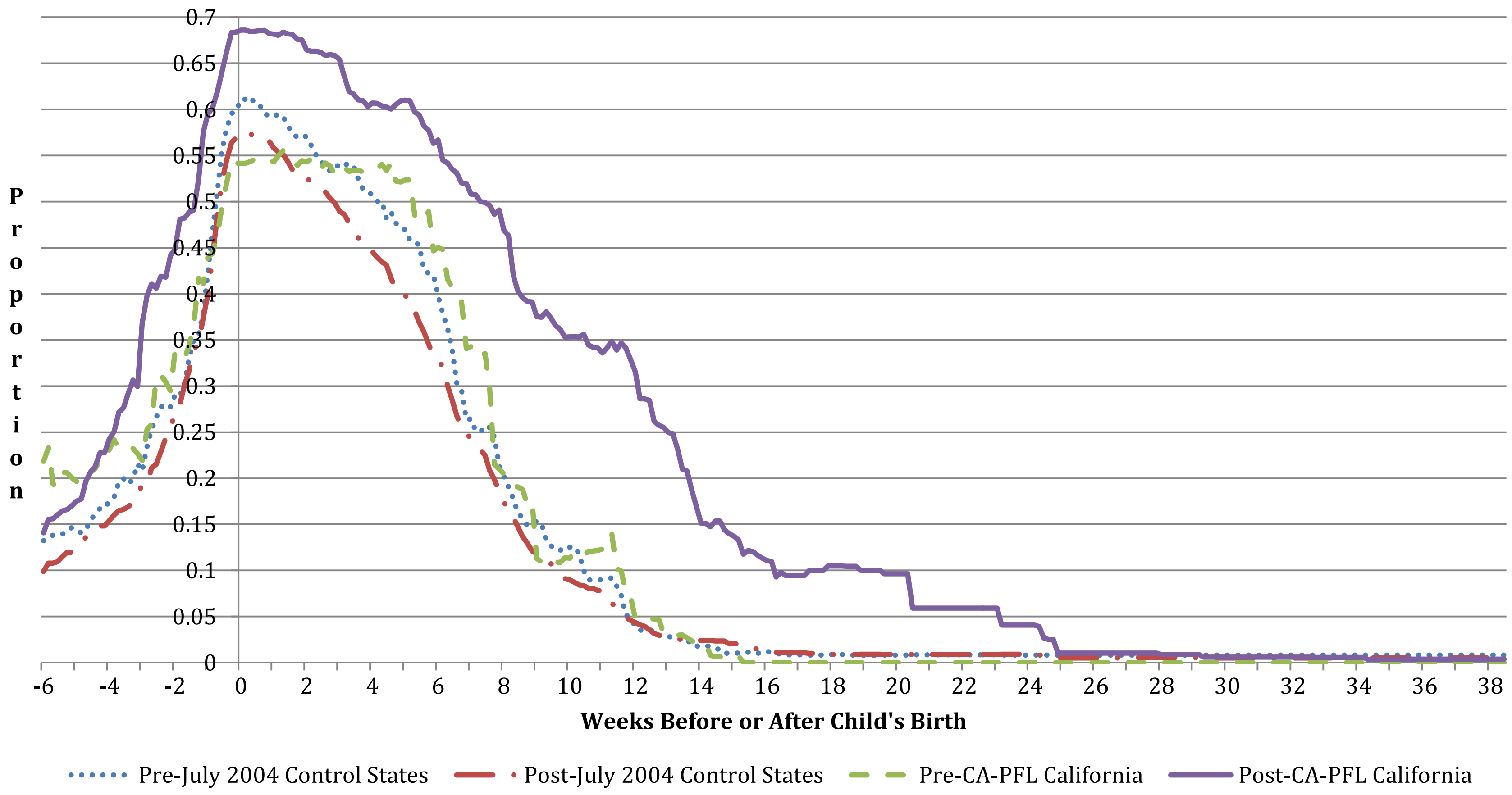
Note: *** $p<0.01$, ** $p<0.05$, * $p<0.1$. Table shows difference-in-difference estimates (coefficients on $CA \times POST$ interactions), with the same controls as in Table 2. Robust standard errors, clustered at the state level, are in parentheses.

Table 6: Regression-Adjusted Estimated Effects of *CA-PFL* on the Probability of Working, Weeks and Hours of Work and Wages for Mothers

	Returned to Work Within One Year Of Birth	Work in 2nd Year After Birth: Annual Weeks Worked	Work in 2nd Year After Birth: Weekly Hours Worked	Log Hourly Wages, One Year After Birth
Mothers Employed at least 32 Pregnancy Weeks from California and Control States				
DD Estimate	0.055***	6.925***	4.209**	0.045
Standard Error	(0.008)	(1.558)	(1.436)	(0.065)
Pre-PFL Baseline	[0.910]	[36.785]	[25.869]	[14.15]
Mothers Employed at all during Pregnancy from California and Control States				
DD Estimate	0.096***	5.803***	3.082**	0.045
Standard Error	(0.013)	(1.303)	(1.099)	(0.044)
Pre-PFL Baseline	[0.813]	[33.067]	[22.866]	[12.61]

Note: *** p<0.01, ** p<0.05, * p<0.1. Table shows difference-in-difference estimates (coefficients on *CA×POST* interactions), with the same controls as in Table 2. Robust standard errors, clustered at the state level, are in parentheses. Pre-PFL sample means in the dependent variables for California mothers are shown in brackets (with levels rather the log of wages displayed in the last column). The probability of working at any job one year after birth is estimated using mothers providing employment information approximately one year (between 47 and 57 weeks) after the birth. Annual weeks and average weekly hours in the second year after birth are measured during the 53rd through 104th weeks after the birth and are not conditional upon employment (i.e. include weeks with zero work hours). Hourly wages, measured in natural logs and 2012-year dollars, refer to those in the first job held during the 47th to 57th weeks subsequent to the birth. Sample sizes are 1,114, 945, 945 and 830 in the top panel for work probabilities, annual weeks worked, weekly hours worked and log hourly wages in the top panel and 1,797, 1,574, 1,574 and 1,208 in the lower panel.

Figure 1: Regression-Adjusted Proportion of Mothers on Leave



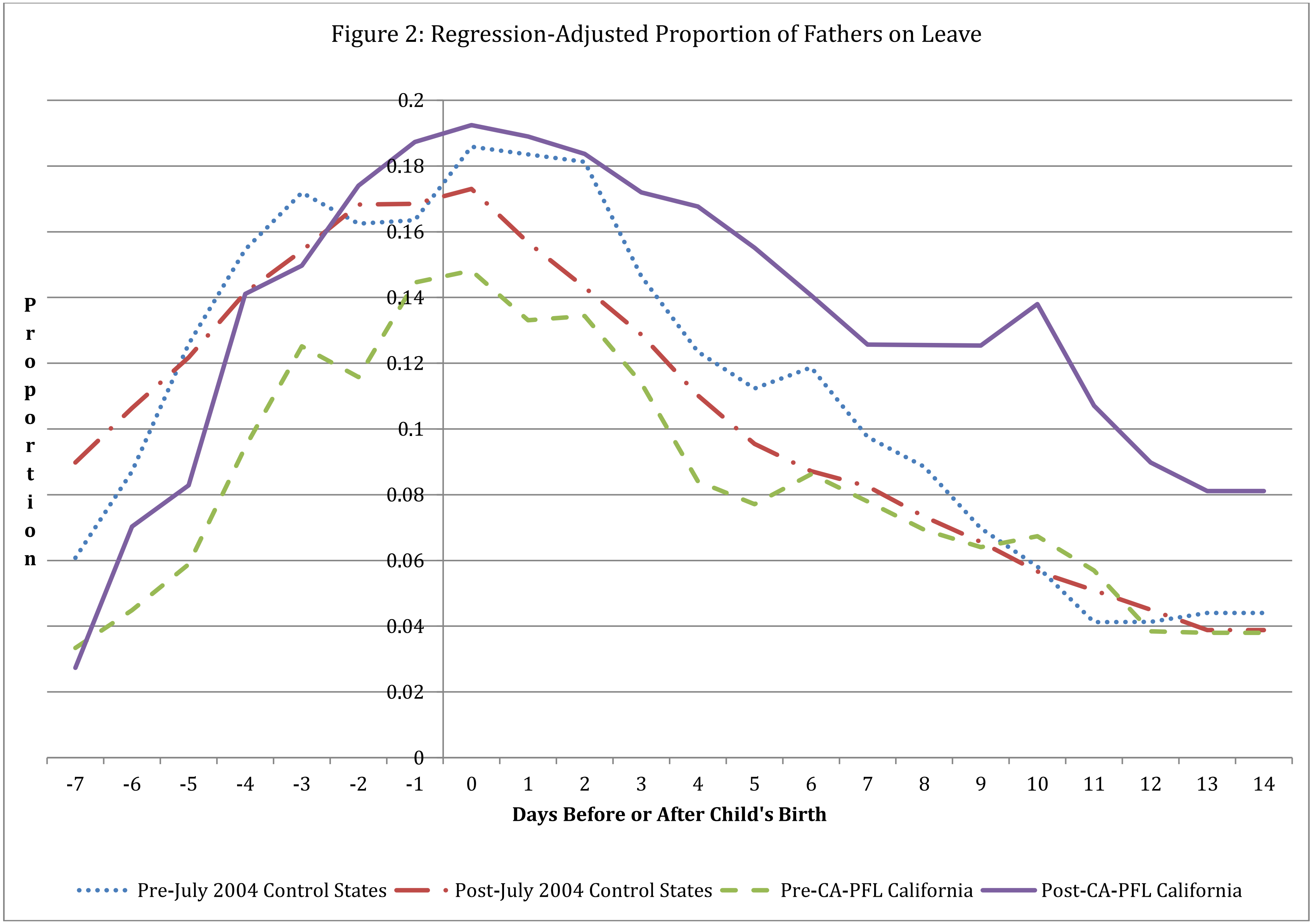


Figure 2: Regression-Adjusted Proportion of Fathers on Leave
Proportion
0.2
0.18
0.16
0.14
0.12
0.1
0.08
0.06
0.04
0.02
0
-7 -6 -5 -4 -3 -2 -1 0 1 2 3 4 5 6 7 8 9 10 11 12 13 14
Days Before or After Child's Birth
Pre-July 2004 Control States
Post-July 2004 Control States
Pre-CA-PFL California
Post-CA-PFL California

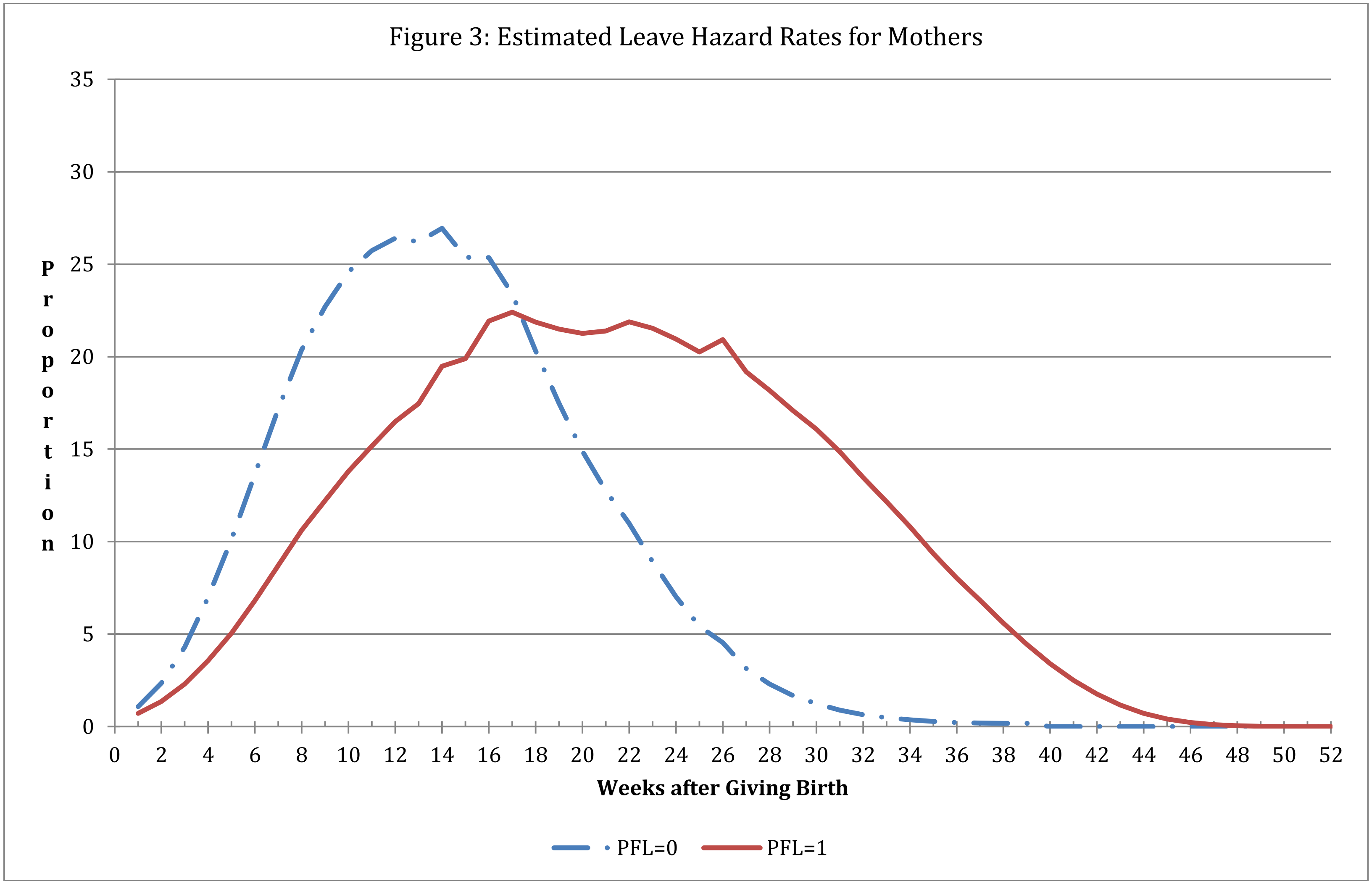


Figure 3: Estimated Leave Hazard Rates for Mothers
Proportion
35
30
25
20
15
10
5
0
0 2 4 6 8 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40 42 44 46 48 50 52
Weeks after Giving Birth
PFL=0
PFL=1

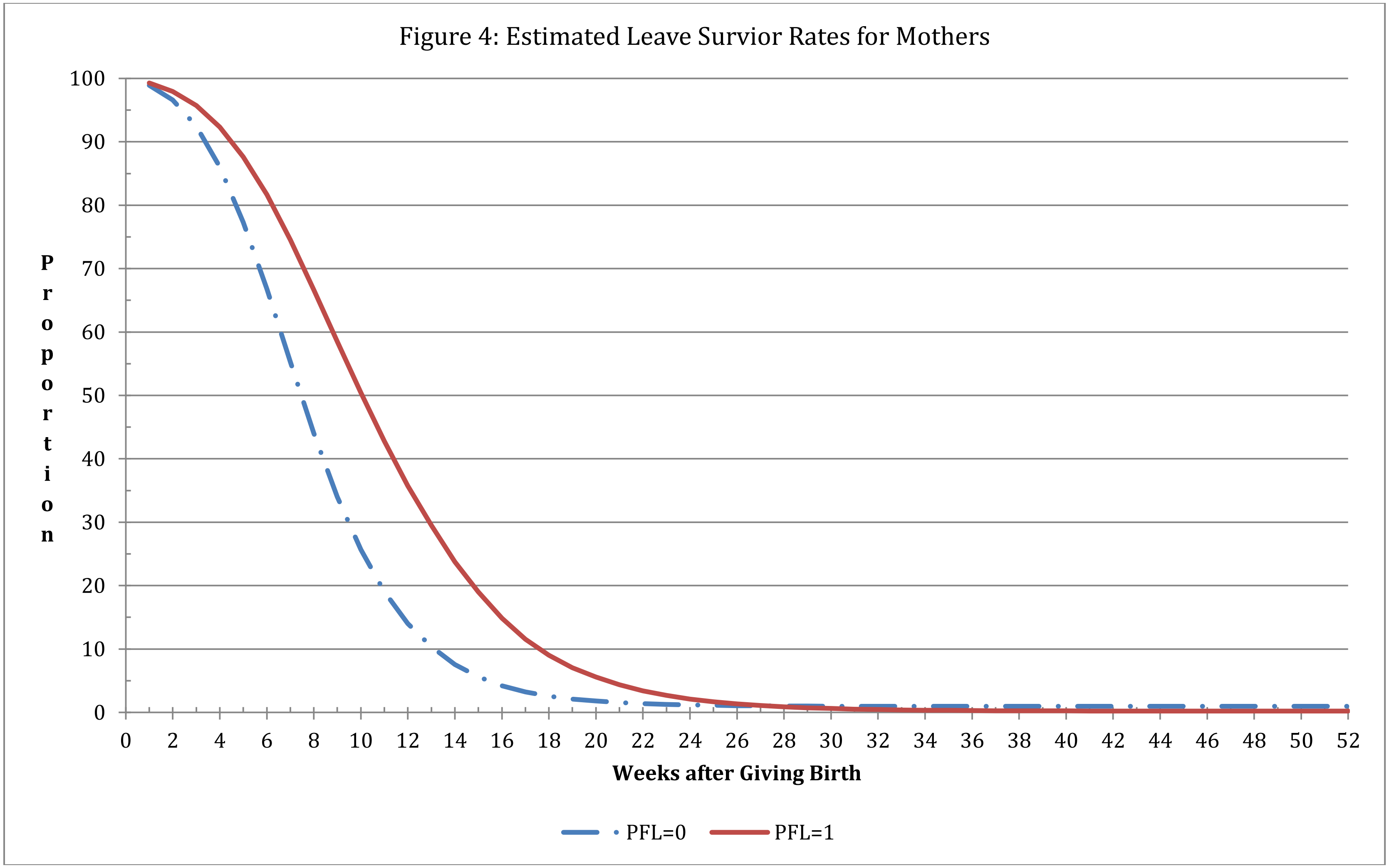


Figure 4: Estimated Leave Survior Rates for Mothers
100
90
80
70
60
50
40
30
20
10
0
Proportion
0 2 4 6 8 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40 42 44 46 48 50 52
Weeks after Giving Birth
PFL=0
PFL=1

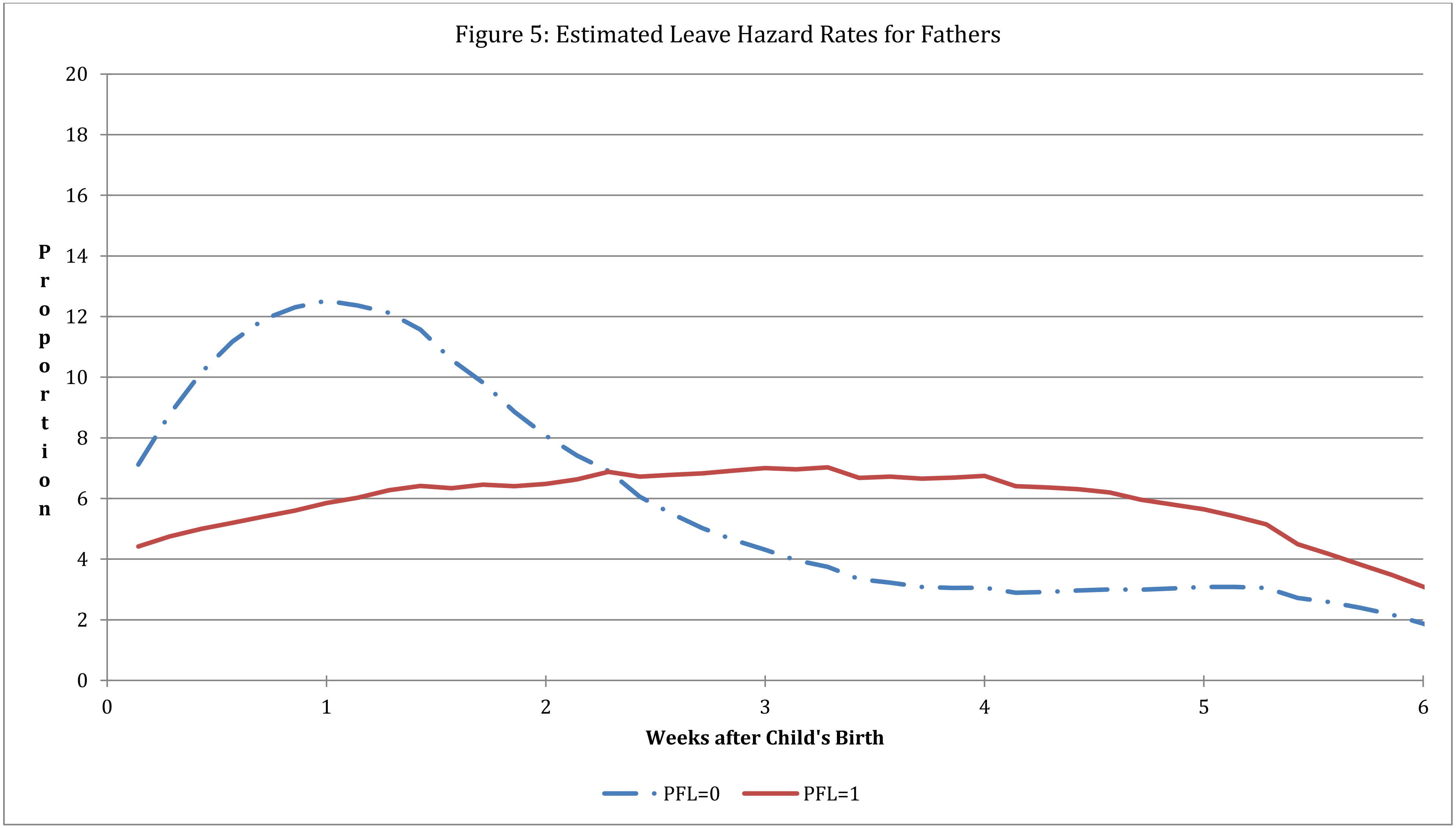


Figure 5: Estimated Leave Hazard Rates for Fathers
Proportion
20
18
16
14
12
10
8
6
4
2
0
0
1
2
3
4
5
6
Weeks after Child's Birth
PFL=0
PFL=1

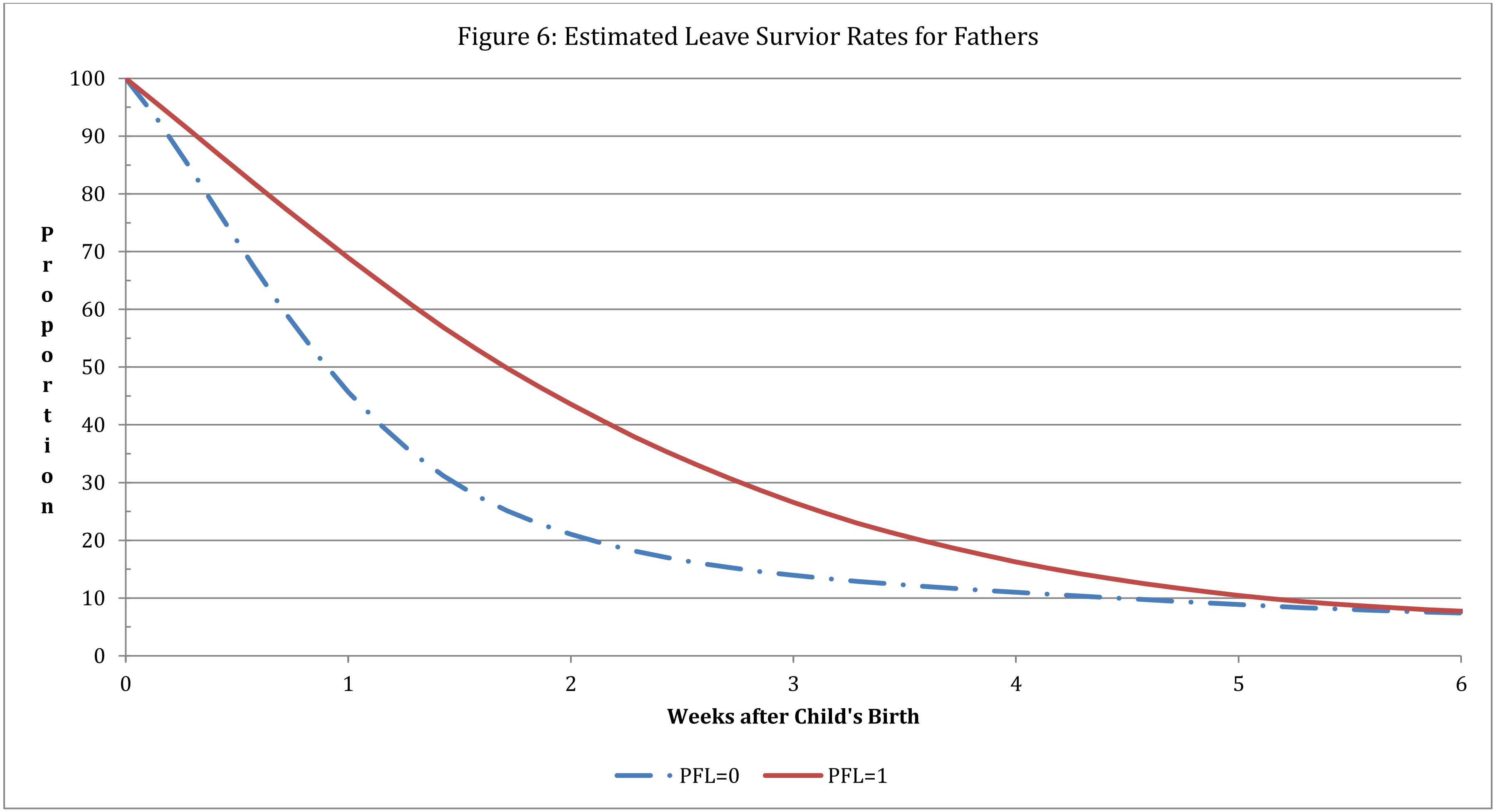


Figure 6: Estimated Leave Survior Rates for Fathers
Proportion
100
90
80
70
60
50
40
30
20
10
0
0
1
2
3
4
5
6
Weeks after Child's Birth
PFL=0
PFL=1

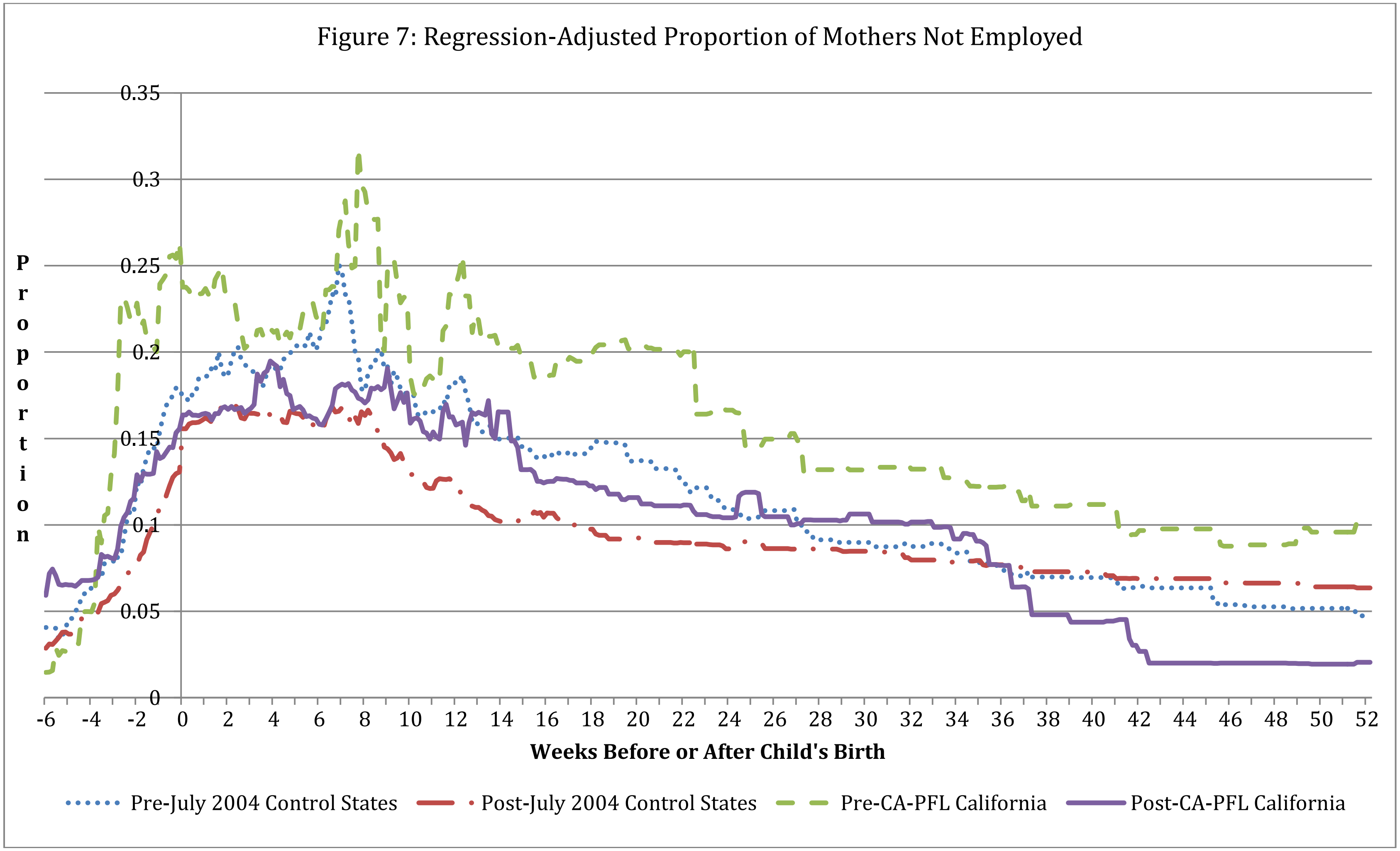


Figure 7: Regression-Adjusted Proportion of Mothers Not Employed
Proportion
0.35
0.3
0.25
0.2
0.15
0.1
0.05
0
-6 -4 -2 0 2 4 6 8 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40 42 44 46 48 50 52
Weeks Before or After Child's Birth
Pre-July 2004 Control States
Post-July 2004 Control States
Pre-CA-PFL California
Post-CA-PFL California

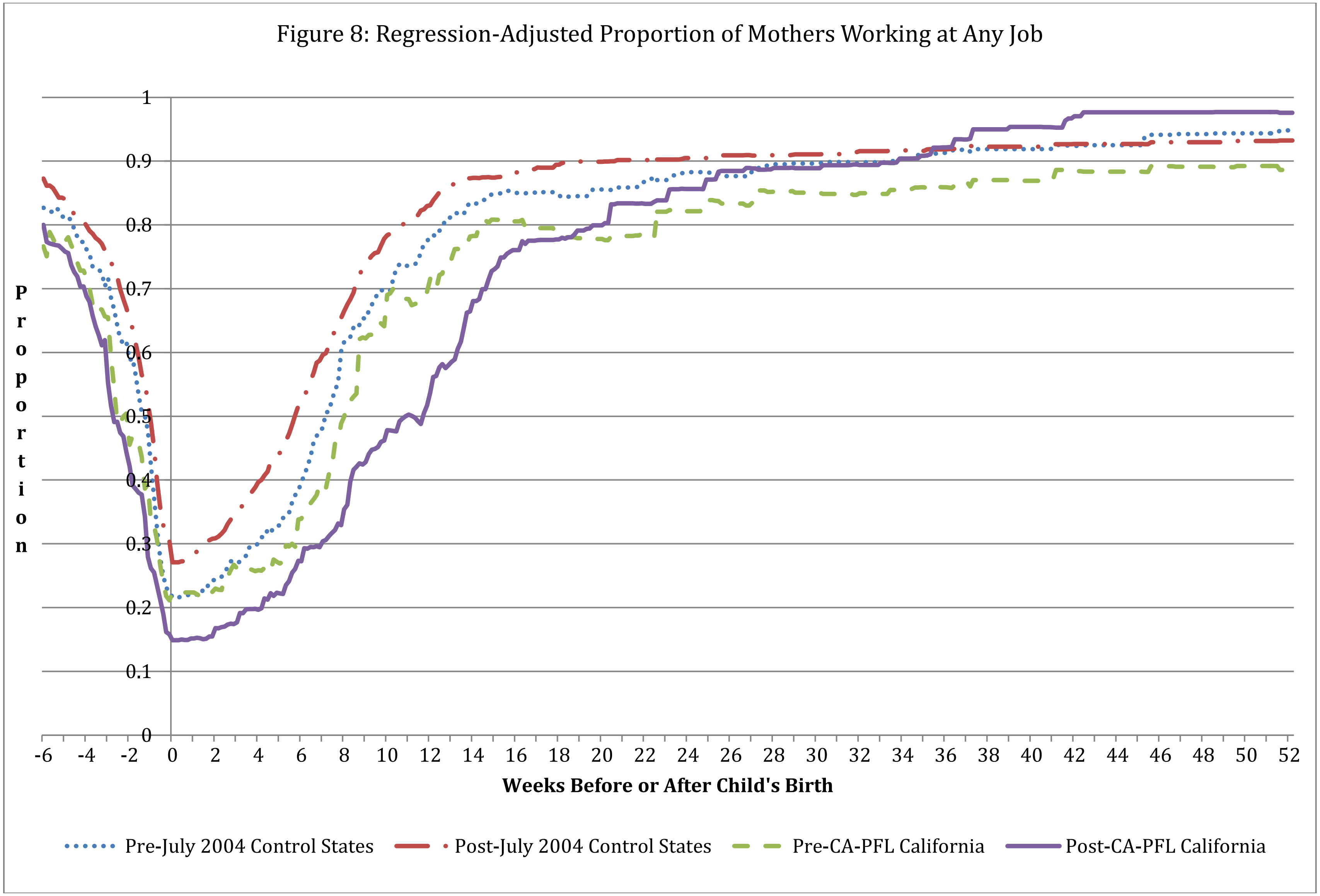


Figure 8: Regression-Adjusted Proportion of Mothers Working at Any Job
1
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
Proportion
-6 -4 -2 0 2 4 6 8 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40 42 44 46 48 50 52
Weeks Before or After Child's Birth
Pre-July 2004 Control States
Post-July 2004 Control States
Pre-CA-PFL California
Post-CA-PFL California

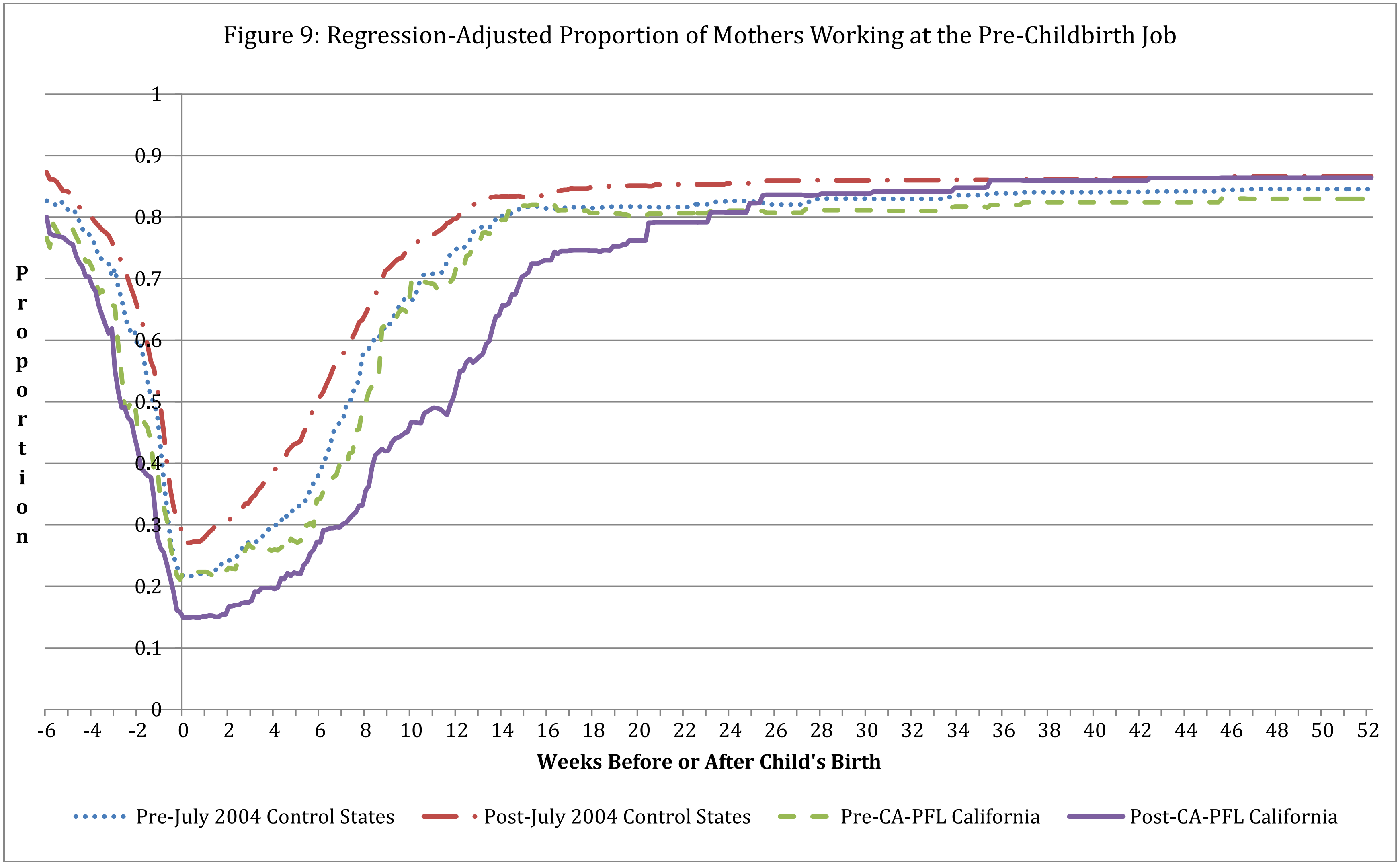


Figure 9: Regression-Adjusted Proportion of Mothers Working at the Pre-Childbirth Job
Proportion
1
0.9
0.8
0.7
0.6
0.5
0.4
0.3
0.2
0.1
0
-6 -4 -2 0 2 4 6 8 10 12 14 16 18 20 22 24 26 28 30 32 34 36 38 40 42 44 46 48 50 52
Weeks Before or After Child's Birth
Pre-July 2004 Control States
Post-July 2004 Control States
Pre-CA-PFL California
Post-CA-PFL California

Table A.1: Descriptive Statistics for Mothers

	Control State Mothers				California Mothers			
	Pre-July 2004		Post-July 2004		Pre-July 2004		Post-July 2004	
Black (=1)	0.272	(0.022)	0.195	(0.013)	0.059	(0.026)	0.055	(0.017)
Hispanic (=1)	0.177	(0.019)	0.164	(0.013)	0.740	(0.048)	0.557	(0.038)
Age (years)	19.621	(0.079)	23.871	(0.085)	19.660	(0.177)	23.855	(0.186)
Education (years)	10.696	(0.084)	12.205	(0.083)	11.352	(0.131)	12.743	(0.145)
Married (=1)	0.157	(0.018)	0.421	(0.017)	0.323	(0.052)	0.445	(0.038)
Experience (/52)	3.177	(0.092)	6.658	(0.099)	2.722	(0.169)	6.468	(0.215)
Family Size	3.662	(0.085)	3.173	(0.051)	4.738	(0.254)	4.090	(0.148)
Child Parity	1.402	(0.033)	1.718	(0.033)	1.317	(0.063)	1.855	(0.079)
Year-2000 Birth	0.101	(0.017)	0.000	(0.000)	0.146	(0.047)	0.000	(0.000)
Year-2001 Birth	0.148	(0.018)	0.000	(0.000)	0.147	(0.039)	0.000	(0.000)
Year-2002 Birth	0.224	(0.021)	0.000	(0.000)	0.292	(0.050)	0.000	(0.000)
Year-2003 Birth	0.308	(0.023)	0.000	(0.000)	0.272	(0.049)	0.000	(0.000)
Year-2004 Birth	0.182	(0.019)	0.078	(0.009)	0.167	(0.041)	0.102	(0.023)
Year-2005 Birth	0.000	(0.000)	0.133	(0.012)	0.000	(0.000)	0.132	(0.026)
Year-2006 Birth	0.000	(0.000)	0.176	(0.013)	0.000	(0.000)	0.125	(0.025)
Year-2007 Birth	0.000	(0.000)	0.199	(0.014)	0.000	(0.000)	0.216	(0.031)
Year-2008 Birth	0.000	(0.000)	0.196	(0.014)	0.000	(0.000)	0.124	(0.025)
Year-2009 Birth	0.000	(0.000)	0.152	(0.012)	0.000	(0.000)	0.225	(0.032)
Year-2010 Birth	0.000	(0.000)	0.067	(0.009)	0.000	(0.000)	0.076	(0.020)

Weighted sample means with standard errors in parentheses. There are 292 pre-July 2004 control mothers, 681 post-July 2004 control mothers, 56 pre-July 2004 California mothers, and 159 post-July 2004 California mothers.

Table A.2: Descriptive Statistics for Fathers

	Control State Fathers				California Fathers			
	Pre-July 2004		Post-July 2004		Pre-July 2004		Post-July 2004	
Black (=1)	0.174	(0.022)	0.171	(0.014)	0.051	(0.030)	0.059	(0.019)
Hispanic (=1)	0.114	(0.019)	0.108	(0.012)	0.609	(0.066)	0.463	(0.040)
Age (years)	19.842	(0.096)	24.066	(0.093)	19.505	(0.217)	24.048	(0.180)
Education (years)	11.561	(0.094)	13.348	(0.096)	11.764	(0.167)	13.177	(0.160)
Married (=1)	0.276	(0.026)	0.501	(0.019)	0.279	(0.060)	0.503	(0.040)
Experience (/52)	3.835	(0.103)	7.220	(0.099)	2.891	(0.195)	6.748	(0.203)
Family Size	3.649	(0.106)	3.060	(0.053)	4.224	(0.254)	3.672	(0.147)
Child Parity	1.487	(0.040)	1.758	(0.036)	1.325	(0.083)	1.769	(0.069)
Year-2000 Birth	0.115	(0.020)	0.000	(0.000)	0.072	(0.036)	0.000	(0.000)
Year-2001 Birth	0.184	(0.023)	0.000	(0.000)	0.153	(0.049)	0.000	(0.000)
Year-2002 Birth	0.259	(0.026)	0.000	(0.000)	0.193	(0.053)	0.000	(0.000)
Year-2003 Birth	0.310	(0.027)	0.000	(0.000)	0.379	(0.065)	0.000	(0.000)
Year-2004 Birth	0.146	(0.021)	0.076	(0.010)	0.129	(0.045)	0.051	(0.017)
Year-2005 Birth	0.000	(0.000)	0.157	(0.014)	0.000	(0.000)	0.181	(0.031)
Year-2006 Birth	0.000	(0.000)	0.172	(0.014)	0.000	(0.000)	0.124	(0.026)
Year-2007 Birth	0.000	(0.000)	0.170	(0.014)	0.000	(0.000)	0.202	(0.032)
Year-2008 Birth	0.000	(0.000)	0.170	(0.014)	0.000	(0.000)	0.174	(0.030)
Year-2009 Birth	0.000	(0.000)	0.181	(0.015)	0.000	(0.000)	0.146	(0.028)
Year-2010 Birth	0.000	(0.000)	0.074	(0.010)	0.000	(0.000)	0.123	(0.026)

Weighted sample means with standard errors in parentheses. There are 246 pre-July 2004 control fathers, 685 post-July 2004 control fathers, 52 pre-July 2004 California fathers, and 143 post-July 2004 California fathers.