Evaluating Workplace Mandates with Flows versus Stocks: An Application to California Paid Family Leave

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Abstract. Employer mandates, as well as other labor demand/supply shocks, are likely to have small or modest wage and employment effects. These effects may be most evident in new hire flows (in number, composition, and wages) rather than in wage and employment levels dominated by incumbents. Using standard household or establishment data, it is difficult to identify small causal effects from shocks. In this paper, we use Quarterly Workforce Indicators (QWI) data to examine the effects of California's paid family leave (CPFL) policy implemented in July 2004, the first state in the U.S. to do so. Among other things, the QWI provides county by quarter by demographic group data on the number and earnings of stable new hires (short-term workers are excluded), the margins over which labor market adjustments are most likely to be evident. The analysis (using double- and triple-differences) shows that CPFL resulted in modestly lower earnings combined with increased employment for young women in California as compared to young men and older women in California, and to younger women, older women, and young men in those states that had paid disability insurance for pregnancy but not paid family leave. Parallel estimates from the Current Population Survey (CPS) are qualitatively similar, but imprecise. Our results are best explained by an outward shift of young women's labor supply, implying that their valuation of paid family leave benefits outweigh their added payroll costs and costs borne by employers due to time-off by experienced employees.

JEL codes: J32 (nonwage labor costs), J38 (public policy)

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1. Introduction

Economic analysis of employer mandates, be they workplace safety, health coverage requirements, family leave policies, or the like, depends crucially on measurement of changes in workplace wages and employment. The costs of mandates is expected to be borne by both employers and employees, with the division of costs a function not only of the relative labor demand and supply elasticities, but also the extent to which workers value mandated benefits. A special case is one in which workers value the expected benefits dollar-for-dollar and the full costs are shifted to workers in the form of lower wages according to their benefit valuation. Under these circumstances, there is no distortion in employment, given that relative labor costs are unchanged, and thus no deadweight welfare loss (Summers 1989, Gruber 1994).

Not surprisingly, economic analyses of workplace mandates typically focus on measuring how wages and employment have been impacted. Because mandates often impact some groups of workers more than others, are implemented in some settings (e.g., states, countries) but not others, and are adopted at different times, evaluation studies most often use differences-in-differences (DD) or triple-difference (DDD) estimators to identify the treatment effects of such policies (e.g., Ruhm 1998; Baum 2003).

This paper examines wage and employment changes following implementation of California's Paid Family Leave (PFL) insurance program in July 2004, the first mandated paid family leave program in the U.S. The theoretical underpinnings and statistical methods used in our analysis are similar to those used in previous studies examining workplace mandates, with one notable difference. Rather than examining changes in the employment and wages among the stock of incumbent employees, we focus on the size and make-up of new hires (employment "flows") and wage offers among new hires. More precisely, we examine changes following enactment of PFL in the wages and number of newly hired young women in California relative to other new hires within California and relative to young women (and other) new hires elsewhere in the country. Data from the Quarterly Workforce Indicators (QWI) (Abowd et al. 2009) are used to measure the wages and employment of "stable" new hires by quarter, location, age, and sex.[1]

Why the focus on new hires? A limitation of existing studies is that the wage and employment effects of workplace mandates do not occur all at once but instead develop gradually over time. We should not expect employers to instantly move to a new equilibrium employment level and/or rapidly change the demographic composition of their workforce following a mandate, although such adjustments are much faster in establishments with normally high rates of separation. Following a mandate employers are unlikely to implement wage cuts for their incumbent workers, cuts that should vary according to how particular workplace groups (say, young women versus others) value a mandated benefit. Although we expect a small impact on incumbent employees (the intensive margin), we should be able to quickly observe treatment effects among new hires (the extensive margin). We expect to see a relative decrease in wage offers for young female new hires and, absent a supply shift, an employment decrease if mandate costs are not fully shifted through lower wages. However, if young women value the benefits and do not bear the full costs, employment may increase.

Even if a workplace mandate has a substantial impact, it is difficult to precisely estimate or even uncover its impact by measuring changes in employment levels and average wages, which are heavily

[1] More standard "stock" (rather than flow) analyses on policy mandates that have differential impacts on wages and employment by state and demographic group are using the Current Population Surveys (CPS). For example, see Card (1992) on minimum wages and Gruber (1994) on health insurance pregnancy coverage. As discussed subsequently, the CPS has been used to examine California's paid family leave by Rossin-Slater et al. (2011), whose principal focus is its effects on time-off from work among new mothers, and by Schroeder (2011, Ch. 3).

weighted by incumbents. If the impact from a workplace mandate is a small or modest effect on new hire employment and wages, these effects are not likely to be discernible in standard measures of employment and wages during the years immediately following implementation.

Although our focus is on employer mandates, specifically, California's paid family leave program, the implications are much broader, applying to any event, behavior, or policy that shifts labor market demand or supply. Labor market adjustments resulting from demand or supply shocks should occur most quickly on the new hire margin, affecting the number and demographic composition of new hires and the wage offers to new hires. Full adjustment to a new long-run equilibrium requires time. Although beyond the scope of this paper, combining new hire information with estimates of incumbent separations and wage growth among incumbents may provide a plausible answer to a difficult question – how long is the long run in labor markets? Other possible applications include the wage and employment effects resulting from immigration, from Hurricane Katrina refugees in destination cities, and from technological change.

2. Overview of California paid family leave policy (CPFL)

Overview/coverage. California's Paid Family Leave (CPFL) policy was enacted August 30, 2002 and took effect July 1, 2004. Prior to the 2004 implementation of CPFL, women had access to paid disability leave during pregnancy and shortly after birth. To understand the marginal effect of California's paid family leave program, one must recognize how it interacts with pre-existing programs and how multiple policies are used in order to receive leave that is job protected and paid. As described below, the principal effect of CPFL has been to extend paid leave among mothers by six weeks.

The California Employment Development Department (EDD) jointly administers the State Disability Insurance (SDI) program, which began in 1977, and the 2004 CPFL. They are jointly financed by a mandatory payroll tax on employees, with no explicit tax on employers. The SDI and CPFL programs provide partial wage replacement for almost all private sector employees. The employees of a business are required to be covered if the business has more than one employee and has paid an employee at least $100 in any quarter during a 12 month reference period. Self-employed and state/local workers are not automatically enrolled, although some can elect coverage. No proof of citizenship is required.

Payroll tax financing. The SDI/CPFL employee tax rate and cap on total contributions have varied substantially across years in order to maintain funds to pay current benefits. As seen in Table 1, the payroll tax rate varied from 0.6% to 1.2% between 2003 and 2011, while the cap on payments varied from a low of $500 in 2007 to $1,120 in 2011.[2] In 2011, the 1.2% employee SDI/CPFL contribution rate combined with a taxable wage ceiling of $93,316 to produce a maximum annual contribution of $1,120. The taxable wage base is adjusted, typically annually, to reflect state wage growth.

SDI wage base and benefit calculation. SDI, as well as the new CPFL program, provides partial wage replacement, with benefits equal to 55% of workers' wages up to a cap. The SDI benefit period is four weeks before the due date and six weeks postpartum for normal pregnancies, but up to eight weeks in the case of Caesarian births or other difficulties (the latter requiring doctor certification). The benefit amount is calculated using a wage base equal to the highest paid quarter during the 12 month reference period 5 to 17 months before the SDI disability claim (eligibility requires at least $300 in earnings during the 12 month reference period). Average SDI pregnancy claim benefits in FY 2011 were $398 per week and the average length of benefits was 10.7 weeks. The 2011 benefit floor was $50 and ceiling $987 per week. CPFL uses the same benefit calculation as does SDI.

[2] In 2003, prior to CPFL, the payment cap was $512. This was increased to $812 in 2004. The cap fell to as low as $500 in 2007 and then rose sharply following the recession, to a high of $1,120 in 2011.

CPFL description. CPFL was created for mothers (or fathers) to bond with their newborns, although it also provides benefits to workers to care for a seriously ill child, spouse, domestic partner, or for a newly adopted child or recently placed foster child.[3] Although California was the first state to provide paid family leave, two others have followed with similarly structured programs.[4] CPFL funds are administered under the SDI umbrella, with employees covered by SDI being eligible for CPFL insurance. Following receipt of six to eight post-partum weeks under SDI, a new mother is eligible for up to six additional weeks of paid family leave using the same benefit formula described above for SDI. In FY 2011, the average CPFL payout was $488 a week for 5.3 weeks. Approximately two-thirds of women receiving SDI pregnancy benefits transition to CPFL benefits.

Job protection vs. paid leave. Although providing partial pay replacement, neither SDI nor CPFL provides job protection. Job protection is provided by state and federal laws guaranteeing unpaid leave.[5] Taken together, a combination of the state SDI and CPFL programs plus the federal FMLA provides a "package" of protected leave with partial wage replacement. And of course some employers provide paid maternity leave independent of any legal requirements. The most generous mandated package includes up to 28 weeks of job protection (up to 16 weeks of pregnancy disability covered by the state PDL concurrent with FMLA, plus 12 weeks protection from the CFRA postpartum) and 16-18 weeks of partial wage replacement (4 weeks pregnancy and 6-8 weeks post-partum under SDI, plus 6 weeks from CPFL).[6]

The transition rate from SDI pregnancy benefits to CPFL claims is well below 100%, being 65.5% in FY 2011. This can occur for several reasons. Some women may prefer or feel a financial need to return to work in order to receive full rather than partial pay. One might expect this to be disproportionately from women in low income households who cannot easily bear the reduced income, highly paid workers whose CPFL benefits are well below 55% of their usual pay due to the benefits cap, and workers for whom promotion and earnings growth is highly dependent on a timely return to work. In addition, workers in small companies (fewer than fifty employees) do not receive job protection through the FMLA and thus may risk losing their job with a lengthy maternity leave. Even absent risk of job loss, a new mother may choose to return to her job at a small company if her employer is highly dependent on her contribution. Not only do some mothers choose not to transition from SDI pregnancy benefits to CPFL, a substantial number take-up CPFL without having taken SDI benefits (reasons for this are not clear; in some cases it may reflect company-provided time off during pregnancy and/or postpartum).

In short, the principal effect of California's 2004 Paid Family Leave program was to extend paid maternity leave by six weeks. Although this is a substantive expansion of benefits, the policy did not

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[3] In FY 2011, 87.3% of CPFL claims were for care of newborns.

[4] New Jersey passed PFL in May 2008, began collecting taxes in January 2009, and began disbursements in July 2009. Washington passed a PFL bill in May 2007, planning to begin payouts in 2012. This has been pushed back, with current plans to begin October 2015.

[5] The 1978 amendments to California's State Fair Employment Practices Act addresses pregnancy discrimination and offers up to four months unpaid, job-protected leave for pregnancy-related disabilities. Pregnancy disability leave (PDL) specifically stipulates that the pregnancy must be a disability and cause the mother to be unable to work (either full or part time). A doctor's note is required and the duration of the leave is up to the doctor. No benefits are paid and the period of leave ends with the birth of the child. Unpaid leave to care for a child following birth is covered by the California Family Rights Act (CFRA), which went into effect in 1992, which provides 12 weeks of unpaid, job-protected leave for private sector employees who have worked the previous 12 months for at least 1,250 hours. Establishments with fewer than 50 employees within a 75 mile radius of the worksite are exempt. The Family Medical Leave Act (FMLA) was signed into federal law a year later with similar provisions and exclusions. Unlike CFRA however, FMLA can be taken both during pregnancy and after the child's birth.

[6] There are additional restrictions for how benefits can be utilized. For example, under some circumstances, the FMLA must be used concurrently during the PDL protected disability.

involve a shift from no mandated paid leave to its current level. Given the incremental nature of the program, identifying CPFL's impact using standard methods and data is likely to prove difficult. Shifting the focus from the policy impact on total wage and employment levels to its impact on new hire wages and employment should provide a more informative approach.

3. *Analysis of employer mandates*

The costs of CPFL are nominally borne by employees through the payroll tax. The costs are attached to all employees, although marginal and average costs per hour are lower for employees with earnings above the tax threshold (about $93 thousand in 2011, an amount likely to be exceeded by relatively few younger workers). The payroll tax costs from CPFL are independent of whether a worker is likely to use and/or values paid family leave. Because the payroll tax is levied at nearly all California establishments, labor supply is highly inelastic (although perhaps more so for men than women), and thus cannot be readily shifted to employers (and/or consumers) for output whose prices are determined nationally or internationally.

Apart from the payroll costs paid by workers, employers face a cost of leave resulting from time off the job among employees taking family leave. Time off among experienced employees reduces output and/or requires added employment by others. An additional cost to employers stems from some degree of uncertainty as to whether a worker on family leave will return. Such uncertainty existed prior to CPFL, but up to six weeks of added leave increases its length. These time-off costs will vary across workplaces. The employer-based costs will attach disproportionately to young female employees, while having far more limited effects on older female and young male employees.[7] Older men will be least affected but, arguably, provide a less attractive control group than do groups similar to young women in either age (young men) or gender (older women) or both (young women in states without PFL).

The expected wage and employment effects resulting from CPFL can be evaluated using the demand and supply "tax incidence" approach (Summers 1989). Effectively, any costs can be thought of as placing a "tax wedge" between labor demand (the total cost to employers) and the labor supply (the net wage to workers). To the extent that labor supply is more inelastic than labor demand, most of the costs are shifted to employees. The payroll costs to employees can be treated as upward labor supply shifts. The "time-off" costs to employers can be represented by a downward shift in demand.

Were there only costs but not benefits from family leave, the employer based costs from time off would lead to lower equilibrium wages and decreased employment among young women, the group most likely to take paid family leave. If labor supply were perfectly inelastic, the payroll tax would be fully shifted to workers, leaving the paid wage constant but lowering the net wage following the tax. To the extent labor supply is not fully inelastic, there will be some increase in the before-tax wage (i.e., some shifting to employers). Costs to employers due to time off by experienced employees are likely to produce lower relative wages and employment for young women (i.e., the principal treated group), with the effects weighted toward a wage rather than employment adjustment the more inelastic is labor supply.

To the extent that employees value the benefit of paid family leave, we would see an outward shift of the labor supply curve. If wages are flexible and can be adjusted downward for those who most value

[7] Young women least likely to have children may place low value on the benefit while bearing costs in terms of job and wage offers due to statistical discrimination. Because employers cannot precisely predict which potential young female hires will have children (or how many), it is reasonable to use all young female new hires as the treatment group. Among young women ages 19-34, the youngest identifiable subgroup among these (ages 19-21 in our data) may be the group for whom employers have the most difficulty discerning likely use of family leave.

the benefits, labor supply shifts out and mitigates or erases negative employment effects as costs are shifted (i.e., little or no deadweight loss). For young women, on average, it is fair to assume that valuation of the benefits exceeds the costs given that most payroll tax receipts that fund family leave are received from male and older female employees. Thus, it is possible that the outward supply shift for young women is sufficiently large to lead not only to a lower wage, but also an increase in employment.

With sufficiently rich data, wage and employment changes for treated versus untreated workers should be evident looking at changes in wage and employment levels. Based on theory, the expectation is that as a result of CPFL, relative wages for the treated (younger women) should decrease relative to other workers, while employment is likely to change little or increase, assuming that the costs can be shifted to young women in the form of lower wages and that these employees value PFL benefits.[8] We expect that such results are most likely to be evident by focusing on data on new hires and new hire wages, the margin over which employers can most easily make adjustments.

4. *Data description: The Quarterly Workforce Indicators*

The employment flow variables at the heart of our analysis are obtained from the Quarterly Workforce Indicators (QWI) database. The QWI is publicly available and is derived from the Local Employment Dynamics (LED) data program, which in turn is built on the confidential Longitudinal Employer-Household Dynamics (LEHD) program. The LEHD is based on state unemployment insurance data and contains individual level quarterly earnings data that matches workers to firms. Crucially for our analysis, the LEHD identifies when a worker begins at a new firm as well as their earnings. The data rely on state participation and while all states have now signed on to participate, eight have not shared data prior to 2000.[9] The QWI provides employment and earnings measures at the state, metropolitan statistical area (MSA), and county levels. Based on individual level LEHD data, these measures are aggregated into narrowly-defined demographic categories including age, sex, ethnicity, race and education within the geographic area. The data cover 98% of all private, non-agriculture employment in the states for which data are available.[10]

Given the incremental nature of California's policy, there is no expectation that wage and employment changes should be substantial. Hence, standard sources such as CPS household data are unlikely to reliably identify wage and employment effects of CPFL (see Rossin-Slater et al. 2011; Schroeder 2011). Careful analysis of the QWI data, however, may enable us to reliably reveal small or modest effects from the policy.

In the analysis that follows, we examine the average number of new hires and the average monthly earnings for these new hires within tightly defined sex-age groupings.[11] All data are observed quarterly. In results shown, we use data for 2002:3 through 2004:2 as the pre-CPFL period and 2004:3 through 2006:2 as the post-treatment period. Thus we have the same number and composition of quarters before and after implementation of the law in July 2004. Examination of the data suggested no apparent effect of the

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[8] Throughout the paper we use the term "young women" and the "treated" synonymously, although about a quarter of paid family leave taking for bonding with children is among men (see Table 1). We do not have data on how duration of PFL differs among male and female recipients.

[9] Data for California is available starting in 1991. In the analysis that follows, five states are excluded from analyses including "all" states. Massachusetts provided no data during our period of analysis. Data for Arkansas, Arizona, New Hampshire, and Mississippi were provided for some but not all quarters.

[10] For a full description of the QWI and its production, see Abowd et al. (2008).

[11] The age groupings identified in the QWI are 14-18, 19-21, 22-24, 25-34, 35-44, 45-54, 55-64 and 65-99. We do not use QWI cells by education or race since many cell sizes would be tiny and suppressed. Education and race change little over our four year period, while state and county fixed effects account for cross-sectional differences.

policy between its passage and eventual implementation in July 2004. We were reluctant to reach back to earlier years because the "tech bubble" had substantial effects through 2001, in particular on the earnings and employment of young men in California, with relatively smaller effects on young women and those outside California.[12]

The unit of analysis is at the demographic-region-quarter level where demographic groups are defined by sex-age group categories and "region" is at the state or county level (results reported in this draft highlight state-level data before turning to county-level analysis). These data allow us to measure average monthly earnings of new employees for the first full quarter in which they are employed. We are able to distinguish all new hires from all new "stable" hires, where a stable hire is defined as an employee who works for at least three consecutive quarters at the firm where they were hired.[13] Our analysis includes employment and earnings data only for stable hires. Among other things, the focus on stable hires avoids our measuring the hiring of temporary replacement workers at non-representative wages.

The narrowly defined demographic and geographic groupings over time in the QWI are ideally suited to help identify treatment effects from California's paid family leave policy. If CPFL affects employment and earnings, then we expect this to be most evident in relative new hire employment and new hire earnings among young women in California. The QWI panel allows us to examine changes that occurred following CPFL among young female treatment groups in California, as compared to changes for other demographic groups within California, as well as compared to young women and other demographic groups outside California.

In order to provide some feel for the QWI data, Table 2 shows average new hire monthly earnings and employment for four demographic groups, young (ages 19-34) women and men and older (ages 35+) women and men, in California, in the four states that have paid SDI (state disability insurance) for pregnancy, and in all states other than California and the five states without complete QWI data during these years (Arkansas, Arizona, Massachusetts, New Hampshire, and Mississippi). We show average earnings for new hires and the number of new hires for the eight quarters before and eight quarters after the July 1, 2004 implementation of CPFL, plus the changes in log earnings and the log of new hires.

Focusing first on the change in log earnings among new hires, we see that both in California and in other states, new hire earnings among young women grew more slowly than for other groups. For example, in California, the increase in earnings was 9 percent, as compared to 10 percent for young men and 11 percent for older women and men.[14] New stable hires among young women in California increased by 9 percent between the two periods, as compared to 7 percent for young men and 6 percent for older women and men. Our statistical analysis will compare changes in young women's new hire earnings and employment following CPFL, as compared to those for young men and older women in

[12] Having said this, our basic results are relatively insensitive to extensions in the treatment and control periods or to omitting data for the quarters immediately before and after implementation.
[13] QWI data are reported with a lag in order that stable hires can be identified retrospectively. In the most narrowly defined groupings, the QWI suppresses data in order to maintain confidentiality. State level data are never suppressed for the sex-age categories. Suppressed county level sex-age data cells are simply dropped. A natural use of the QWI is to use it to estimate employment and earnings levels ("stocks") as well as new hire flows. Levels data are inappropriate for this analysis, however, since a quarterly payroll can include both women going on leave and replacement workers who may be added.
[14] Earnings are in nominal dollars; inflation is accounted for in regressions using quarter fixed effects. In the paper we refer to the change in the log of mean earnings as the percentage change. It measures a percentage change in earnings with an intermediate base in the denominator and has the advantage of being invariant to the base. Of course, the difference in the log of mean earnings is not the same as the difference in the means of log earnings.

California, and compared to young women (and others) outside California. The results of our subsequent analysis, which indicate slightly lower earnings and higher employment for young women due to CPFL, can be gleaned to a limited degree from the information in Table 2.

5. *Method of analysis*

As implied by the summary statistics in Table 2, there are three major sources of variation that are exploited to identify the impact of CPFL on new hires and new hire earnings. We begin by setting up a simple difference-in-differences (DD) model that uses only the demographic variation within California to identify the impact on hires and wages. Then we progress to a model that includes data from other states, thus utilizing geographic variation to identify estimated treatment effects, but focusing on a comparison group of young women outside of California.

Consider the following simple econometric specification, which will serve as a basis for our analysis of the labor market impacts of the CPFL.

$$\ln(Y_{dq}) = \beta_T \ Post_q \text{ x } Young\_Fem_d \ + \delta_d + \gamma_q + \epsilon_{dq} \qquad (1)$$

In this specification only data from California is used. The unit of observation is at the demographic-quarter level with Y_{dq} representing either total new hires or the average monthly earnings of those new hires in a given demographic group (*d*), in a given quarter (*q*). The key coefficient of interest is β_T, which measures the impact on young-female new hires or earnings following implementation of CPFL. The variable *Post* is an indicator variable equal to one for all observations in or after the third quarter of 2004, after CPFL went into effect.[15] The variable *Young_Fem* is an indicator variable equal to one for the 19-21, 22-24 and 25-34 categories.[16] The variables δ_d and γ_q represent full sets of demographic group and quarter indicator variables in order to control for time invariant differences between demographic groups as well as shocks that may have hit all demographic groups in a given quarter. A richer variant of this model adding greater variation to the data is also estimated using as the observation unit California counties by quarter rather than state by quarter (this specification includes county fixed effects).

Equation 2 presents a DD model that expands the data to include other states, but restricts the comparison group and sample to observations for young women.[17] Including other states (or counties from other states) allows us to directly compare changes to hiring and wage offers for young women in California with young women in other states not impacted by CPFL.

$$\ln(Y_{sq}) = \beta_T \ Post_q \text{ x } CA_s \ + \gamma_q + \alpha_s + \epsilon_{sq} \qquad (2)$$

In equation (2), the treatment effect estimate β_T represents log differences in new hires and new hire earnings for young women in California following CPFL as compared to outcomes for young women throughout the rest of the country (or in those four SDI states with paid pregnancy but not paid family leave benefits), conditioned on fixed effects for quarter *q* and state *s*. Equation (3) is also estimated using

[15] In preliminary analysis, we failed to find a separate passage effect.

[16] These are the age groupings that are most likely to be impacted by the CPFL. Birth per 1,000 women in 2004 were 20.1 for 15-17 year olds; 66.2, 96.3, 110.5, and 97.7 for age groups 18-19, 20-24, 25-29, 30-34 (close to our ages 19-34 treatment group); and 46.5 and 10.1 for women 35-39 and 40-44 (Martin et al. 2011, Table 4). We initially provide estimates based on a combined *Young_Fem* group of women ages 19-34, but we can also provide separate estimates for the three “treated” age groups identified in the QWI, ages 19-21, 22-24, and 25-34.

[17] We provide estimates based both on use of all other states as controls and just the four SDI states (Hawaii, New Jersey, New York, and Rhode Island) whose disability programs provide partial wage replacement benefits for pregnancy, but not paid family leave, as was the case in California prior to implementation of PFL in 2004.

the county as the unit of observation (substituting county for state fixed effects) for a comparison of California counties with counties in the four SDI states.

An alternative to equation (2) is to extract the same estimated treatment effect from a more general triple-diff model that includes all states (or all counties across all states) and all demographic groups, but still identifies β_T off the comparison of employment and wage changes for young women in California compared to young women in other states. It takes the form

$$\ln(Y_{sdq}) = \beta_T \ Post_q \text{ x } CA_s \text{ x } Young\_Female_d \ + \delta_{sd} + \gamma_{sq} + \alpha_{dq} + \epsilon_{sdq}, \qquad (3)$$

where the variables δ_{sd}, γ_{sq} and α_{dq} represent full sets of state-demographic group, state-quarter and demographic group-quarter indicator variables in order to control for time invariant differences between state-demographic groups as well as shocks to demographic groups and states that occur in a given year. In what follows, we estimate the triple-diff models using both state-quarter and county-quarter observations.

The inclusion of these large sets of indicator variables effectively controls for many of the worker differences that vary across demographic groups, states (or counties), and years. Consider education, a crucial determinant of new hire earnings. If young women in California have different levels of education than other demographic-state or county combinations these differences will be picked up by δ_{sd} as long as they are time invariant over this period. Furthermore, if state education levels or demographic group education levels are changing over time these changes will be picked up by γ_{sq} and α_{dq} respectively.

County rather than state level results will naturally provide greater variation to the outcome variables of interest than will the state and are likely to provide more precise estimates. There are two (minor) disadvantages of the county-level data. First, these models become quite large given the substantial number of interaction variables that are required in fixed effects models. Second, county-level data provide somewhat greater noise than do state data. Indeed, the QWI does not report data for very small data cells in order to insure confidentiality (this involves a tiny proportion of total county-by-demographic observations).

6. *Estimates of CPFL treatment effects on new hire earnings and employment*

Tables 3-5 provide estimates of the "treatment" effects of the California PFL on the earnings and employment of young (ages 19-34) new hires. Table 3 provides a simple comparison within California in the quarters prior to and following implementation of CPFL. The left side of the table compares earnings and employment for young women compared to young men, while the right-side compares young women to older women. Using state-level observations, two specifications are shown, one absent fixed effects (to which we attach little weight) and one with fixed effects for quarter and detailed age groups and race. Columns (3) and (5) show the fixed effects model using county observations. Not surprisingly, the combination of modest causal effects from CPFL combined with a relatively small number of observations leads to generally insignificant coefficients in the state-level analyses. Most estimates using county-level observations are significant (standard errors are clustered).

The within-California comparison indicates that, as expected, CPFL leads to lower wage offers to young female new hires, on the order of about 1 percent as compared to young men and 2 percent compared to older women. There is no evidence of decreased hiring of young women. The estimates consistently show higher rates of hiring of about 2 percent relative to young men and 1-2 percent compared to older women. Taken together, the results indicate that paid family leave is valued by young women, leading to an increase in labor supply that allows employers to shift CPFL costs at least partially to young women and increase employment offers. Because nominal payroll taxes are borne by men and older women as well as young women, we cannot rule out that such results using the within-California comparison occur in part from decreased labor supply and higher wages among the comparison groups.

In Table 4, CPFL wage and employment effects are estimated through a comparison of treated young women in California compared to young women in other states. The left-side presents a comparison with young women in "all" states (more precisely, 44 control group states). Our preferred comparison is between young women in California and those in the four SDI states, like California, that have paid disability benefits for pregnancy but did not implement additional paid family leave as did California in 2004. This narrower comparison should prove a more accurate measure of the wage and employment effects resulting from CPFL's addition of six paid weeks of family leave on top of existing paid disability benefits.

The Table 4 results using the SDI comparison states are qualitatively similar to those seen in Table 3 using the within California comparison. New hire earnings among young women in California following CPFL fall by about a half to one percent relative to young female new hire earnings in the four SDI states, a somewhat lower wage decrease estimate than seen in the within-California analysis. As seen previously, the hiring of young women increases in California relative to the SDI comparison states by just under 2 percent. Similar analysis using an all-state comparison is less clear-cut, with small positive earnings effects. None of the coefficient estimates in Table 4 are statistically significant at standard levels.

In Table 5, we move from the simple double-diff estimates presented above to a triple-diff comparison across states, demographic group, and time (before and after treatment). Using the SDI states as the control markets, we compare new hire earnings and employment young women in California as compared to young men and older women in the SDI states. We get a pattern of results similar to those shown previously, each suggesting small negative effects of CPFL on new hire wages for young women, coupled with small increases in new hire employment. Using SDI state young men as the comparison group, earnings for young California women falls by about ½ percent and new hire employment rises by 1 percent. Compared to older women, earnings fall by about 1 percent and employment rises by 1 percent. Estimates are similar when including fixed effects for quarter, state, and demographic group (columns 2 and 5) and when adding a full set of interaction fixed effects between quarter, state, and demographic group.

Table 6 shows results using an identical triple-diff framework as did the results in Table 5, but instead uses the county-by-demographic-by-quarter observations rather than state-demographic-quarter. Estimates with county level data are substantially more precise and somewhat larger in magnitude. As compared to both young men and older women in SDI comparison states, the new hire earnings for California young women fell by 2 percent (with results highly significant), while employment of young stable female new hires increased by over 1 percent.

Overall, we are impressed by the relative consistency and plausibility of the estimates in Tables 3-6, despite the fact that statistical significance is sometimes marginal.[18] It is reasonable to conclude, at least preliminarily, that California's PFL has led to slightly lower relative wages for young women, coupled with slightly higher rates of new hire employment offers. The results are consistent with an increase in

[18] Two caveats are in order. First, a future version will include falsification tests that can determine if estimates of this magnitude and significance might arise by chance. Second, in analyses not shown, we estimate separate treatment effects for women ages 19-21, 22-24, and 25-34. Negative wage effects were much larger for the 19-21 group than for the older groups, a surprising result we plan to explore further.

labor supply among young women due to CPFL benefits and an efficient shifting (at least partially) of CPFL costs to young women through lower wages. For the within-California comparison, we cannot rule out that the results partially reflect small labor supply decreases among men and older women due to incomplete cost shifting (i.e., their increased payroll taxes are not fully offset by wage increases). It is likely that such effects for California young men and older women are limited, however, since we obtain highly similar results based on comparisons across states, where the comparison groups of men and older women are unaffected by California payroll taxes.

7. *Employment and earnings treatment effect estimates from the Current Population Survey*

Although California's paid family leave (CPFL) program is relatively recent, we are aware of three studies (unpublished) that analyze, at least briefly, the employment and earnings effects of CPFL using household data from the Current Population Survey. Not surprisingly, estimates of the effects have been small and imprecise, making it difficult for authors to make reliable inferences as to program effects. Two studies have used the March Annual Social and Economic Supplement (ASEC) to the CPS, and one study (on which our analysis below builds) has used the Monthly Outgoing Rotation Groups (MORGs) of the CPS. The latter provides samples roughly three times larger than the March surveys and reports earnings and hours on the principal job the week prior to the survey, rather than annual earnings, weeks, and hours the previous calendar year.

In a paper analyzing the FMLA, Espinola-Arrendondo and Mondal (2010) also examine CPFL employment effects using the March 2001-2007 CPS. They compare female employment changes in California following CPFL relative to changes for women in other states with and without expanded FMLA provisions. Using numerous combinations of treatment and comparison groups, the authors conclude that all their treatment estimates are "both economically and statistically insignificant." One possibility is that that true effects of CPFL are close to zero. But another is that short-run CPFL effects are most likely to show up in data on new hires and not among incumbent employees. It is not surprising that CPFL effects (which may well be small) cannot be discerned using small samples in the March CPS to examine changes in employment and earnings *levels* largely determined by incumbent employees.

Although not the principal focus of their paper, Rossin-Slater et al. (2011) also use the March CPS to examine the earnings and employment effects of CPFL.[19] The authors faced several difficulties in precisely identifying those who are treated and not treated by CPFL (time of a child's birth cannot be reliably identified), so they present several alternative comparisons and make adjustments to their estimates to better measure treatment effects. They conclude that there were no changes in employment following CPFL, but that work hours (hours last week and in the prior year) among treated groups increased, conditional on employment. The authors offer a possible explanation for these somewhat puzzling results, but note that such an explanation is speculative and that future study is needed.

For reasons outlined previously, our preferred approach is to analyze the CPFL (as well as other labor market shocks) using data on new hires. For purposes of comparison, however, we also provide analysis using data from the Monthly Outgoing Rotation Groups (MORGs) of the Current Population Survey (CPS) for 2001-2009 (the year 2004 is dropped). In this draft, we summarize results previously presented with these data in Schroeder (2011). In our next version of the paper, we will provide CPS-

[19] Rossin-Slater et al. (2011) provide evidence that CPFL significantly increased time off from work among mothers of young children. This is the primary focus of their paper. If time off from work is sufficiently costly to employers and not fully offset by benefits from greater long-run attachment, then we should observe negative wage effects among young female new hires.

MORG estimates that more directly comparable to those presented previously from the QWI (e.g., using similar treatment and comparison groups to those in the QWI).

The sample includes private sector non-student wage and salary workers (neither government nor self-employed workers are automatically covered by CPFL), ages 18-65. In addition, workers whose earnings are imputed by Census, about 30% of the sample, are omitted. As shown by Bollinger and Hirsch (2006), it is essential that they be excluded. Non-respondents are assigned earnings from "similar" donors, but match criteria do not include state, sector of employment, marital status, or numerous other variables (the only attribute with an exact match is sex). Thus, a worker in California is likely to be assigned the earnings of a resident outside California in a different industry, among other differences. Inclusion of imputed earners leads to substantial attenuation of coefficients (so-called "match bias") on non-match criteria, in this case treatment effect estimates based on residence in California.[20] The final sample consists of almost 1.7 million observations, about half women. Sample sizes for treatment groups of young women in California are obviously far smaller.

Wage and employment equations are estimated using a triple-diff approach, comparing alternative groups of young women in California to older female and to young and old male workers in other states. Our wage variable measures average hourly earnings, calculated in one of four ways depending on the employee, 1) hourly straight-time wage for those who are paid by the hour and do not have tips, overtime or commission earnings, 2) the straight-time wage plus weekly tips, overtime or commissions divided by the usual hours worked per week, 3) usual weekly earnings (inclusive of tips, overtime and commissions) divided by the usual number of hours worked for salaried workers, and 4) weekly earnings divided by hours worked the previous week for salaried workers whose usual hours vary. In the CPS, employment is defined as those employed either at work or absent from work. Individuals on PFL should be counted as employed with their wages, usual weekly earnings, and usual weekly hours reported. Log wage and LPM employment equations are estimated, with controls for potential work experience (defined as age minus years of schooling minus 6, plus its square and cube), the monthly state unemployment rate by state, and dummies for sex (which are also interacted with experience), marital status, presence of child, schooling degree, metropolitan size, race/ethnicity, citizenship status, part-time, occupation, industry, and fixed effects for states and years. PFL treatment effects for young women in California can be compiled based on changes over time in California relative to alternative comparison groups outside California.

Not surprisingly, treatment effect estimates indicate small and imprecise effects on wages and employment. However, almost all estimates of PFL wage effects on California women (relative to alternative groups outside California) are negative, ranging from about 0.5 to 2 percent. Employment effects relative to alternative male groups outside California are effectively zero, but more often positive than negative. No within-California estimates are provided. Estimated employment effects are mostly positive, but estimated imprecisely and generally small, mostly between 0.5 and 2 percent. Falsification tests (using pseudo-laws in other states) produced estimates that were very small, never statistically significant, and that displayed no qualitative pattern.

At this point we withhold judgment on the CPS analysis, waiting instead to align the CPS treatment and control groups to better correspond to our new hire analysis using the QWI. That said, the preliminary

[20] It is relatively simple to reweight the estimation sample of respondents to make it representative using inverse probability weights (IPW) with respect to response. In practice, estimates based on unweighted respondent samples and IPW samples produce highly similar results (Bollinger and Hirsch 2006). Retaining imputed earners, while introducing serious match bias, fails to account for non-ignorable response bias since imputed earnings are from respondent donors (non-respondent earnings cannot be observed).

evidence, while imprecise, supports evidence from the QWI indicating small wage penalties and positive employment effects for young women in California resulting from the introduction of paid family leave.

8. *Conclusion*

Employer mandates are likely to have small effects. Nonwage benefits highly valued by workers relative to their costs are those most likely to be voluntarily provided by employers (with costs shifted to workers). Benefits that have substantial costs relative to worker valuation are those least likely to be mandated through the political process. Mandated worker benefits not provided voluntarily but that are politically viable are likely to have small or similar benefits and costs (Addison and Hirsch 1997).

Unfortunately, most available data sets are incapable of accurately identifying small or modest causal effects from employer mandates. Specifically, household data sets such as the CPS have small sample sizes of individuals by geographic location by time period. Establishment data rarely provide the demographic and geographic breakdown needed to analyze mandates that differentially impact alternative groups of workers. More fundamentally, wages and employment across demographic groups or within businesses change gradually. Incumbent workers are not likely to have their pay reduced substantially. Nor will businesses quickly alter the demographic make-up of their trained workforces through dismissals. The margin for which one is most likely to observe wage and employment adjustments in response to an employer mandate is with respect to new hires, both through changes in their demographic composition and in the wages offered.

The Quarterly Workforce Indicators (QWI) data set provides a relatively new and underutilized resource that lends itself well to the evaluation of public policies that differentially affect employment and/or earnings with respect to time, location, and demographic group.[21] Particularly appealing is QWI's provision of data on the number and earnings of stable (not short-term) new hires, the margins over which labor market adjustments are most likely to occur.

In this paper, we use the QWI to examine the effects of California's paid family leave policy adopted in July 2004, the first state in the U.S. to do so. CPFL effectively added six weeks of partially paid leave to new mothers (or fathers), added to the ten to twelve week paid disability leave already available for pregnancy and the postpartum period. Recent work by Rossin-Slater (2011) indicates that CPFL led to increased time off among mothers with infants. Our preliminary analysis concludes that CPFL resulted in modestly lower earnings combined with increased employment for young women in California, as compared to young men and older women in California and to younger women, older women, and young men in those states that had paid disability insurance for pregnancy but not paid family leave.

As a check on the plausibility of our estimates, it is possible to produce a "back-of-the-envelope" guesstimate of the type of wage adjustment that would occur if there were full shifting of costs to young women. To produce such an estimate, we need information on the program costs (the explicit tax cost for CPFL plus additional workplace costs), the use of paid family leave by newly-hired employees (the number of and average length of leaves) and their expected tenure, the growth in real wages, and a discount rate. Our initial foray into this exercise produced estimates of wage effects on the order of 1 to 2 percent, surprisingly similar to our empirical estimates. The next version of our paper will provide such estimates formally and with refinements in our method and assumptions.

[21] An excellent example is a recent paper by Gittings and Schmutte (2012), who use the QWI to examine the effect of minimum wages on new hires and separations.

Our results are best explained by an outward shift of young women's labor supply, implying that their valuation of paid family leave benefits outweigh their added payroll costs and costs borne by employers due to time-off by experienced employees. Stated alternatively, at least some of the costs of CPFL have been shifted to those workers who most highly value the benefit (young women), leading to no employment or efficiency loss. Had we observed only within-California evidence, an equally plausible explanation might be that costs were only partially shifted to young women and that the observed pattern also reflects lower take-home earnings (after CPFL worker payroll taxes) and lower employment among men and older women. This possibility is unlikely to be important, however, given that we find similar evidence using cross-state comparisons where the worker control groups are unaffected by California payroll taxes.

References

Abowd, J.M., B.E. Stephens, L. Vilhuber, F. Andersson, K.L. McKinney, M. Roemer, and S. Woodcock. 2009. "The LEHD Infrastructure Files and the Creation of the Quarterly Workforce Indicators," in T. Dunne, J.B. Jensen and M.J. Roberts, eds., *Producer Dynamics: New Evidence from Micro Data* (Chicago: University of Chicago Press), 149-230.

Addison, John T. and Barry T. Hirsch. 1997. "The Economic Effects of Employment Regulation: What Are the Limits?" in B.E. Kaufman, ed. *Government Regulation of the Employment Relationship*, IRRA 50th Anniversary Volume, Madison, WI: Industrial Relations Research Association, 125-178.

Baum, Charles L., II. 2003. "The Effects of Maternity Leave Legislation on Mothers' Labor Supply after Childbirth." *Southern Economic Journal* 69(4) April, 772-799.

Bollinger, Christopher R. and Barry T. Hirsch. 2006. "Match Bias from Earnings Imputation in the Current Population Survey: The Case of Imperfect Matching." *Journal of Labor Economics* 24(3) July, 483-519.

Card, David. 1992. "Regional Variation in Wages to Measure the Effects of the Federal Minimum Wage," *Industrial and Labor Relations Review* 46(1), October, 22-37.

Espinola-Arrendondo, Ana and Sunita Mondal. "The Effect of Parental Leave on Female Employment: Evidence from State Policies." Washington State University School of Economic Sciences working paper series WP 2008-15 (2008). Revised 2010.

Fass, Sarah. 2009. *Paid Leave in the States: A Critical Support for Low-wage Workers and Their Families*, National Center for Children in Poverty, Columbia University, March.

Gittings, R. Kaj and Ian Schmutte. 2012. "Getting Handcuffs on an Octopus: Minimum Wages, Employment and Turnover," unpublished manuscript, Louisiana State Univesity and he University of Georgia.

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

Martin, J.A., B.E. Hamilton, S.J. Ventura, M.J.K. Osterman, S. Kirmeyer, T.J. Mathews, and E.C. Wilson. 2011. *Births: Final Data for 2009*, National Vital Statistics Reports, Division of Vital Statistics, U.S. Department of Health and Human Services. Volume 60, Number 1, November 3, 2011.

Rossin-Slater, Maya, Christopher J. Ruhm, and Jane Waldfogel. 2011, "The Effects of California's Paid Family Leave Program on Mothers' Leave-Taking and Subsequent Labor Market Outcomes," NBER Working Paper 17715, December.

Ruhm, Christopher J. 1998. "The Economic Consequences of Parental Leave Mandates: Lessons from Europe." *Quarterly Journal of Economics* 113(23) February, 285-316.

Schroeder, Mary C. 2011. *The Economics of Mandated Paid Leave*, doctoral dissertation, Department of Economics, Emory University.

Summers, Lawrence H. 1989. "Some Simple Economics of Mandated Benefits," *American Economic Review Papers and Proceedings* 79(2) May, 177-183.

Table 1: Descriptive Statistics on California Paid State Disability Insurance (SDI) and Paid Family Leave (PFL)

SDI/PFL claims and benefits	FY 2005	FY 2006	FY 2007	FY 2008	FY 2009	FY 2010	FY 2011
Total SDI pregnancy claims paid	172,623	175,194	183,013	189,139	181,685	169,957	168,593
SDI claims transitioning to PFL bonding claims			108,818	115,392	119,442	111,024	127,529
Estimated PFL/SDI share			0.655	0.631	0.636	0.614	0.655
Average weekly benefit, SDI pregnancy claims			$354	$368	$382	397	$398
Average weeks, SDI pregnancy claims			11.97	10.43	10.43	10.50	10.70
Average weekly benefit, PFL claims	$409	$432	$441	$457	$472	$488	$488
Average weeks per PFL claim	4.84	5.32	5.37	5.35	5.39	5.37	5.30
Total PFL claims filed	150,514	160,988	174,838	192,494	197,638	190,743	204,893
Total PFL claims paid	139,593	153,446	165,967	182,834	187,889	180,675	194,777
Total PFL benefits paid*	$300.42	$349.33	$387.88	$439.49	$472.11	$468.79	$498.44
% of PFL claims filed for bonding	87.7%	87.8%	87.6%	87.6%	88.8%	87.8%	87.3%
number of bonding claims filed by women	109,566	112,631	119,893	129,986	132,958	123,632	128,774
% of bonding claims filed by women	83.0%	79.7%	78.3%	77.1%	75.8%	73.8%	72.0%
CY SDI/PFL tax, contribution, benefit rules	**CY 2000**	**CY 2001**	**CY 2002**	**CY 2003**	**CY 2004**	**CY 2005**	**CY 2006**
Contribution rate	0.65%	0.70%	0.90%	0.90%	1.18%	1.08%	0.80%
Taxable wage ceiling	$46,327	$46,327	$46,327	$56,916	$68,829	$79,418	$79,418
Maximum worker contribution	$324	$324	$417	$512	$812	$858	$635
Maximum weekly benefits	$490	$490	$490	$603	$728	$840	$840
	CY 2007	**CY 2008**	**CY 2009**	**CY 2010**	**CY 2011**		
Contribution rate	0.60%	0.80%	1.10%	1.10%	1.20%		
Taxable wage ceiling	$83,389	$86,698	$90,669	$93,316	$93,316		
Maximum worker contribution	$500	$693	$997	$1,026	$1,120		
Maximum weekly benefits	$882	$917	$959	$987	$987		

* dollar amounts are in millions

Source: Data were compiled by authors from data provided on the website and by an analyst at the State of California, Employment Development Department.

Table 2: Descriptive Evidence on QWI New Hire Earnings and Employment, Pre- and Post-CPFL

	Pre-CPFL		Post-CPFL		Post minus Pre	
	Average New Hire Earnings	Average New Hires	Average New Hire Earnings	Average New Hires	ΔlnEarnings	ΔlnNewHires
California						
Young women	$1,562	77,370	$1,706	84,627	0.088	0.090
Young men	$1,875	85,413	$2,068	91,439	0.098	0.068
Older women	$2,111	61,302	$2,359	65,283	0.111	0.063
Older men	$3,371	72,684	$3,758	77,050	0.109	0.058
Paid SDI states						
Young women	$1,486	15,431	$1,620	16,517	0.087	0.068
Young men	$1,846	15,349	$2,037	16,242	0.098	0.057
Older women	$2,037	12,630	$2,252	13,322	0.100	0.053
Older men	$3,518	13,991	$3,893	14,723	0.101	0.051
"All" states except CA						
Young women	$1,311	13,350	$1,410	14,479	0.073	0.081
Young men	$1,714	13,979	$1,882	15,117	0.093	0.078
Older women	$1,751	10,674	$1,920	11,627	0.092	0.086
Older men	$3,053	11,777	$3,343	12,863	0.091	0.088

Earnings and new hires are monthly values based on quarterly averages. Paid SDI states (i.e., those with state disability insurance covering pregnancy and postpartum) are Hawaii, New Jersey, New York, and Rhode Island. Because of incomplete data in the QWI, the "All" states group does not include Arkansas, Arizona, Massachusetts, New Hampshire, and Mississippi. They include the four paid SDI states. Young women and men are ages 19-34.

Table 3: CPFL Effects on New Hire Earnings and Employment: Diff-in-Diff within California with Young Men and Young Women Comparison Groups

Control Group:	Young Men			Older Women		
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	New Hire Earnings			New Hire Earnings		
Post x Female 19-34	-0.0056	-0.0098	-0.0134***	-0.0275**	-0.0226	-.0226***
	(0.0150)	(0.0152)	(0.0049)	(0.0136)	(0.0149)	(0.0058)
Observations	96	96	5358	96	96	5367
R-squared	0.114	0.997	0.959	0.291	0.996	0.951
FE	None	Qtr, Dem	Qtr, Dem, County	None	Qtr, Dem	Qtr, Dem, County
Dependent variable:	New Hires			New Hires		
Post x Female 19-34	.0295	0.0231	0.0236*	0.0267	0.0156	0.0130
	(0.0215)	(0.0319)	(0.0130)	(0.0265)	(0.0312)	(0.0122)
Observations	96	96	5415	96	96	5414
R-squared	0.032	0.994	0.994	0.149	0.993	0.995
FE	None	Qtr, Dem	Qtr, Dem, County	None	Qtr, Dem	Qtr, Dem, County

*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors in parentheses. In cols. 1, 2, 4, and 5, s.e. are clustered at the demographic level (12 sex by age dummies); cols. 3 and 6 at the county by demographic level. The number of county fixed effects differs across samples due to non-disclosure of small cells.

Table 4: CPFL Effects on New Hire Earnings and Employment:
Diff-in-Diff across States with Young Women Comparison Group

Control Group:	All States		SDI States		
	(1)	(2)	(3)	(4)	(5)
Dependent variable:	NH Earnings	NH Earnings	NH Earnings	NH Earnings	NH Earnings
Post x California	0.0091	0.0096	-0.0052	-0.0062	-.0083
	(0.0087)	(0.0079)	(0.0042)	(0.0042)	(0.0142)
Observations	2160	2160	240	240	7105
R-squared	0.038	0.186	0.019	0.058	0.942
FE	None	Qtr, Dem, State	None	Qtr, Dem, State	Qtr, Dem, County
Dependent variable:	New Hires	New Hires	New Hires	New Hires	New Hires
Post x California	0.0022	0.0009	0.0211	0.01592	0.0183
	(0.0139)	(.0151)	(0.0146)	(0.0127)	(0.0147)
Observations	2160	2160	240	240	7132
R-squared	0.274	0.858	0.417	0.964	0.997
FE	None	Qtr, Dem, State	None	Qtr, Dem, State	Qtr, Dem, County

*** p<0.01, ** p<0.05, * p<0.1. Demographic fixed effects are for detailed age groups among young women. Robust standard errors in parentheses. In cols. 1-4, s.e. are clustered at the state by demographic level and in col. 5 at the county by demographic level. The number of county fixed effects differs across samples due to non-disclosure of small cells. The SDI states are Hawaii, New Jersey, New York, and Rhode Island. Excluded from the "All States" regressions are Arkansas, Arizona, Massachusetts, New Hampshire, and Mississippi. "All States" estimates at the county level are not included due to length of time required for estimation with county by demographic fixed effects.

Table 5: CPFL Effects on New Hire Earnings and Employment:
Triple Diff among SDI States with Young Men and Older Women Comparison Groups

Control Group:	Young Men /SDI States			Older Women/SDI States		
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	NH Earnings	NH Earnings	NH Earnings	NH Earnings	NH Earnings	NH Earnings
Post x Calif. x Female 19-34	-0.00284	-0.00497	-0.00307	-0.0182	-0.0117	-0.00854*
	(0.0183)	(0.0184)	(0.00468)	(0.0179)	(0.0181)	(0.00438)
Observations	480	480	480	480	480	480
R-squared	0.112	0.978	0.998	0.252	0.976	0.998
Dependent variable:	New Hires	New Hires	New Hires	New Hires	New Hires	New Hires
Post x Calif. x Female 19-34	0.00850	0.0101	0.0111	0.000319	0.00983	0.00984
	(0.0296)	(0.0315)	(0.0126)	(0.0300)	(0.0302)	(0.0106)
Observations	480	480	480	480	480	480
R-squared	0.434	0.995	1.000	0.435	0.992	1.000
FE	None	Qtr,State,Dem	State-Qtr Dem-Qtr State-Dem	None	Qtr,State,Dem	State-Qtr Dem-Qtr State-Dem

*** p<0.01, ** p<0.05, * p<0.1. Robust standard errors in parentheses. In cols. 1-6, s.e. are clustered at the state by demographic (sex by age) level. The SDI states are Hawaii, New Jersey, New York, and Rhode Island.

Table 6: CPFL Effects on New Hire Earnings and Employment with County Fixed Effects: Triple Diff among SDI States with Young Men and Older Women Comparison Groups

Control Group:		Young Men /SDI States			Older Women/SDI States	
	(1)	(2)	(3)	(4)	(5)	(6)
Dependent variable:	NH Earnings	NH Earnings	NH Earnings	NH Earnings	NH Earnings	NH Earnings
Post x Calif. x Female 19-34	-0.0221	-0.0234*	-0.0227***	-0.0352**	-0.0218*	-0.0198***
	(0.0138)	(0.0116)	(0.00641)	(0.0135)	(0.0112)	(0.00618)
Observations	14,222	14,222	14,225	14,225	14,225	14,225
R-squared	0.085	0.891	0.969	0.148	0.846	0.958
Dependent variable:	New Hires	New Hires	New Hires	New Hires	New Hires	New Hires
Post x Calif. x Female 19-34	0.00288	0.00444	0.0136	-0.0264	0.00385	0.0144*
	(0.0375)	(0.0319)	(0.00892)	(0.0326)	(0.0296)	(0.00762)
Observations	14,222	14,222	14,276	14,225	14,225	14,276
R-squared	0.119	0.848	0.998	0.128	0.842	0.998
FE	None	Qtr, Dem, County	County-Qtr Dem-Qtr County-Dem	None	Qtr, Dem, County	County-Qtr Dem-Qtr County-Dem

*** p<0.01, ** p<0.05, * p<0.1. Table 6 is structured the same as Table 5, but with the units of observation being counties rather than states. Robust standard errors in parentheses. In cols. 1-6, s.e. are clustered at the county by demographic (sex by age) level. The number of county fixed effects differs across samples due to non-disclosure of small cells.