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The Impact of the Family and Medical Leave Act

Jane Waldfogel

Abstract

This article uses data from employer surveys and the March Current Population Survey to investigate the impact of the Family and Medical Leave Act (FMLA) on coverage, leave-taking, employment, and earnings. The variation in state laws prior to the FMLA and the variation in coverage under the FMLA provides a "natural experiment" in which the effect of the law can be compared for treatment and control groups. Although the FMLA covers less than half of workers in the private sector (many of whom already had coverage pre-FMLA), this article finds that leave coverage and usage did increase post-FMLA. The other surprising finding is that this mandated benefit had no significant negative effects on women's employment or wages. ©1999 by the Association for Public Policy Analysis and Management.

INTRODUCTION

This article examines the impact of the 1993 Family and Medical Leave Act (FMLA) on coverage, leave-taking, employment, and earnings. The FMLA has been the subject of a good deal of controversy, yet there has been little research on it to date. Research on this topic is particularly important in light of the recent trends in employment for women, as 54 percent of new mothers are now back at work within a year of their child's birth [U.S. Bureau of the Census, 1993]. Such research is also important in the wake of state and federal welfare reforms that emphasize moving women from welfare to work; these reforms increasingly affect women with young children.

Women consistently report that family and medical leave policies are important in helping them manage their work and family responsibilities, yet women, and particularly low-income women, are less likely than men to be covered by employer policies in this area [Commission on Family and Medical Leave, 1996]. Critics of the FMLA contend that it does little to help those who lack coverage because it covers less than half of all private sector workers and because those who are covered are

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disproportionately likely to have been covered already. The FMLA has also been criticized for imposing costs on employers, which will in turn be passed on to women employees. This study provides the first empirical evidence on both these questions. It finds that the FMLA did in fact lead to increased coverage and usage for some working women, without imposing significant costs in terms of lower employment or wages for women overall.

BACKGROUND

Although the FMLA includes several types of family and medical leave for both men and women, it is best known as the first federal law to provide maternity leave rights for (some) American women. Prior to the FMLA, the United States was the only industrialized country without a federal law guaranteeing a job-protected maternity leave [Kamerman and Kahn, 1995].¹ The FMLA, which had been vetoed twice by President Bush, was the first piece of legislation passed under the new Clinton administration in February 1993. The FMLA requires employers with 50 or more workers to offer a job-protected family or medical leave of up to 12 weeks to qualifying employees (those who have worked at least 1250 hours in the prior year) who need to be absent from work for reasons that meet the terms of the law (for example, an employee's own illness including maternity-related disability, or the need to care for a newborn or an ill family member). The law does not require paid leave, but it does require employers who provide health insurance coverage to continue to do so during the leave period.

The FMLA went into effect in August 1993, and its impact on leave policies varies according to state. In the absence of federal legislation, many states had passed state leave laws prior to 1993.² Of particular interest in this analysis, 11 states and the District of Columbia (shown in the Appendix) had laws in place that required firms to offer a job-protected maternity leave a year or more before the FMLA [Irwin and Silberman, 1993; Waldfogel, 1994; Women's Legal Defense Fund, 1993].³ Firm-specific characteristics matter, too. Large firms and unionized firms were particularly likely to have maternity leave policies in place prior to the passage of the FMLA [Hyland, 1990; Spalter-Roth and Hartmann, 1990].

The first part of this article looks at the effect of the FMLA on coverage, taking into account the fact that many of those workers covered by the law were already covered by state laws or by preexisting provisions in union contracts or employer policies. This part of the study also investigates the extent to which leave coverage varies by gender.

¹ The Pregnancy Discrimination Act of 1978 required employers with disability plans to cover pregnancy, but it did not provide coverage for employees whose firms did not offer such plans [Blau, Ferber, and Winkler, 1997].

² According to the Commission on Family and Medical Leave [1996], a total of 34 states had passed laws prior to the FMLA. Of these, 11 covered state employees only. The other 23 covered both private and public sector workers, but with varying provisions.

³ A 12th state, Vermont, had a state law that went into effect pre-FMLA (July 1992) but after the date of the first pre-FMLA observations used here (March 1992); for that reason, I code it as a state that had no state law pre-FMLA. As noted earlier, a total of 34 states had some leave legislation prior to the passage of the FMLA, but many of these states' laws were not fully comparable to the FMLA. For example, 11 states covered state employees only, while 11 other states covered public and private sector employees but did not provide job-protected maternity leave [Waldfogel, 1998b].

The second part of the article looks at the impact of the law on the utilization of leave, particularly among women with children under the age of one. Prior research by Klerman and Leibowitz [1997] provides some evidence that state maternity leave laws led to increased leave-taking over the 1980s, although their results varied by specification.⁴ I extend that research by looking at the effects of the FMLA on leave-taking. As noted earlier, the likely magnitude of these effects is unclear since the law does not cover all working women and because many of those whom it does cover were likely to have been covered already.

The third and fourth parts of the article examine the impact of the law on affected groups' employment and earnings. The family leave provisions have been estimated to cost employers approximately \$250 per year for each employee using the leave.⁵ As noted earlier, although the FMLA applies to both men and women, only women use maternity leave, and they may be more likely to use the other leave benefits as well because women typically have more responsibility than men for other family members' care. Theory on mandated benefits [Burtless, 1995; Summers, 1989] predicts that the costs of the leave provisions might therefore be borne by women as a group, in the form of lower employment (as employers might shift away from hiring women) and or in the form of lower wages (as employers might pass on to women the costs of the benefit). However, these effects might be offset by an increase in the labor force attachment of individual women. Recent research on the effects of maternity leave provisions across 16 European countries [Ruhm, 1998] found that the effects varied by the length of the leave mandated; short mandated leaves had positive employment effects and no wage effects, while long mandated leaves had negative effects on both employment and wages.⁶ There has been no research to date on the employment and earnings effects of the FMLA. Research by Gruber [1994] on a related but more expensive mandate (health insurance coverage for maternity, which was estimated to cost employers nearly \$1000) found a fall in both employment and wages.⁷ Prior research, then, suggests that the FMLA may have negative employment or earnings effects, but that these are likely to be small because the leave mandated by the FMLA is short and because it is relatively low in cost.

RESEARCH DESIGN

The variation in the impact of the law by state of residence creates a "natural experiment," in which the effect of the law can be compared by assessing changes pre- and postlaw across groups. Here, I take advantage of the fact that most states (the 39 "experimental" states) had no state law providing the right to a job-protected maternity leave prior to the FMLA whereas some states (the 11 "nonexperimental" states plus the District of Columbia) did have state laws requiring job-protected maternity leave prior to the FMLA. I then use difference-in-difference analysis to compare the change in outcomes for the experimental states to the change in outcomes

⁴ Research on young women in the United States and Britain also suggests that higher coverage leads to increased usage [Waldfoegel, 1997, 1998b].

⁵ This figure represents the average employer's cost (in 1993 dollars) of maintaining health insurance coverage for an employee on leave, taking into account that not all employers provide such coverage [Waldfoegel, 1994]. There may also be other costs associated with leave (for example, the cost of hiring a replacement worker), but there are no good estimates for these.

⁶ See also Ondrich, Spiess, and Yang [1997a, 1997b] who examine the effects of long leaves in Germany.

⁷ Gruber [1994] also found that the fall in participation was offset by an increase in hours worked.

for the nonexperimental states. This methodology is superior to a simple time-series analysis in that it uses the states for whom there was no law change as a control group for the states where there was a law change. However, this approach will produce biased estimates if there were other differences between these two groups of states that affected both the passage of state laws and the change in their outcome variables over time. For instance, states that passed laws might have done so at the same time that their labor market was improving (or worsening); if so, this unobserved heterogeneity could bias the estimated effect of the law on an outcome such as employment upwards (or downwards). A related problem is endogeneity. States may have passed leave legislation because more women were using leave already (or because many women wanted to use leave but were not able to do so). If so, the endogeneity of the state law would bias the estimated difference-in-difference estimate of the effect of a law change on an outcome variable such as leave-taking upwards (or downwards).

To address these potential sources of bias, I use a difference-in-difference-in-difference methodology.⁸ This requires identifying one or more “treatment” groups (those affected by the law) and one or more “control” groups (those not affected). Because the law was expected to mainly benefit women of childbearing age, particularly those with children and most particularly those with newborns, I use women aged 19 to 45 with children, and women aged 19 to 45 with children under the age of one, as my two treatment groups. Childless women aged 19 to 45 and men aged 19 to 45, who may be suitable substitutes for mothers aged 19 to 45 in the labor market but who are expected to be less likely to use leave, constitute the two control groups for the leave-taking analysis; in the analysis of employment and wages, I use older women (aged 46 to 60) as an additional control group, because the law change might affect the employment and wages of childless women if they are of childbearing age (that is, aged 19 to 45). Because firm size also matters in terms of preexisting coverage and coverage under the FMLA, I use the Current Population Survey (CPS) data on number of employees to further differentiate treatment and control groups in the analysis.⁹ Finally, because only those working for an employer and those working more than 1250 hours during the prior year are covered by the FMLA, I exclude self-employed and part-time workers in the analysis of leave-taking and wage effects.¹⁰

It is important to note that the difference-in-difference-in-difference approach can resolve the potential endogeneity bias only if one assumes that the characteristic of the women in the treatment group that was correlated with the passage of the law (for example, their desire to take leave) was also a characteristic of the individuals in the control group. If this assumption does not hold, even difference-in-difference-in-difference estimates could be biased by endogeneity.¹¹

⁸ See Hamermesh [1998] for a discussion of natural experiments as a solution to the endogeneity problem.

⁹ I do not use union status because it is not clear whether unionized employees should be seen as a control group in that they were much more likely to have been covered pre-FMLA [see, for example, Hyland, 1990] and therefore might have a smaller increase in leave usage post-FMLA, or as a treatment group in that they might be more likely to have information about their expanded rights under FMLA [see Commission on Family and Medical Leave, 1996] and therefore might use more leave post-FMLA.

¹⁰ Although the CPS does have data on hours worked over the past year, it does not record hours worked prior to the onset of leave. For that reason, I use the usual hours worked variable to determine whether an employee is usually full-time (35 or more hours per week) or part-time (less than 35 hours per week) and then restrict the analysis to full-time workers only. As noted in the text, the self-employed are also excluded because the FMLA only covers employees.

QUESTION 1: DID COVERAGE INCREASE AS A RESULT OF THE FMLA?

There are two sources of data on this question,¹² and both point to an increase in coverage following the passage of the FMLA.

The Bureau of Labor Statistics' Employee Benefits Survey provides information about coverage at medium and large firms (100 or more employees) in 1988, 1989, 1991, 1993, and 1995; it surveyed firms with fewer than 100 employees in 1990, 1992, and 1994. As shown in Tables 1 and 2 (and Figure 1), there are sharp increases in maternity leave coverage in 1993, 1994, and 1995. The percentage of full-time employees in medium and large firms whose employers provide maternity leave (paid or unpaid) increased from 35 percent in 1988 to about 40 percent in 1989 and 1991 and then jumped to 63 percent in 1993 and 86 percent in 1995;¹³ in smaller firms (defined here as those with under 100 employees), the percentage rose from 19 percent in 1990 to 20 percent in 1992 to 49 percent in 1994. Trends for paternity leave are similar, although less dramatic. Are these increases in coverage due to the FMLA? As noted earlier, state laws, union contracts, and voluntary employer policies have been and continue to be important sources of coverage, but none of these factors can explain the more dramatic rise in coverage beginning in 1993. The FMLA, in contrast, seems a more likely explanation.¹⁴

A second source of data is the 1994 Westat survey of small, medium, and large firms [see Cantor et al., 1995] conducted for the Family and Medical Leave Commission.¹⁵ This survey found that 11 percent of firms, employing 60 percent of all workers, were covered by the FMLA. Overall, an estimated 46 percent of all workers qualified for FMLA coverage (as nearly a quarter of covered workers did not meet the requirement that they have worked at least 1250 hours in the prior year), and women were only slightly less likely to be covered and to qualify than men.¹⁶ The survey asked how many of the covered employers had to change their

¹¹ An alternative approach that would not require this assumption would be to use instrumental variables to predict the passage of the state laws; however, in this instance, it is hard to identify an instrument that would be correlated with passage of the state laws but not with the outcome variables of interest.

¹² The National Longitudinal Survey of Youth (NLSY) is another potential source of data on this question, but only one wave (1993) of post-FMLA data was available at the time that this article was written.

¹³ The fact that coverage did not reach 100 percent in 1995 suggests that at least some firms had not yet come into compliance with the FMLA.

¹⁴ The 1993 Employee Benefit Survey was administered during the period November 1992 through October 1993, so some observations precede the passage of the FMLA (February 1993) and its effective date (August 1993). However, the FMLA was widely anticipated throughout the survey period because of the election in November 1992 of a Democratic president who was expected to promptly sign into law a measure that had previously passed but had been vetoed twice by a Republican president. Thus, it is reasonable to interpret the increase in the 1993 survey as a first indication of the effects of the FMLA.

¹⁵ The FMLA legislation established this bipartisan commission to study and report back to Congress about mandatory and voluntary family and medical leave policies. The Commission's final report was released on 1 May 1996 [see Commission on Family and Medical Leave, 1996].

¹⁶ These figures are nearly identical to those estimated by Bond [1995]; using the Changing Workforce Survey, he projects that 46 percent of all American workers (43 percent of women and 48 percent of men) would qualify under the FMLA. Note that both sets of figures refer to private sector workers only. The FMLA also covers all public sector workers, regardless of firm size, so estimates including public sector workers are higher. For example, the Westat survey found that 66 percent of all public and private sector employers work for covered employers and 55 percent qualify for coverage.

Table 1. Percentage of full-time employees with maternity leave coverage.

Medium and large firms (100 or more employees)	1988	1989	1991	1993	1995
Total with unpaid maternity leave	33	37	37	61	84
Total with paid maternity leave	2	3	2	2	2
Total with any maternity leave	35	40	39	63	86

Small firms (less than 100 employees)	1990	1992	1994
Total with unpaid maternity leave	17	18	47
Total with paid maternity leave	2	2	2
Total with any maternity leave	19	20	49

Source: U.S. Department of Labor, Bureau of Labor Statistics, 1996, 1998.

Table 2. Percentage of full-time employees with paternity leave coverage.

Medium and large firms (100 or more employees)	1988	1989	1991	1993	1995
Total with unpaid paternity leave	16	18	26	54	84
Total with paid paternity leave	1	1	1	0	2
Total with any paternity leave	17	19	27	54	86

Small firms (less than 100 employees)	1990	1992	1994
Total with unpaid paternity leave	8	8	47
Total with paid paternity leave	0	1	2
Total with any paternity leave	8	9	49

Source: U.S. Department of Labor, Bureau of Labor Statistics, 1996, 1998.

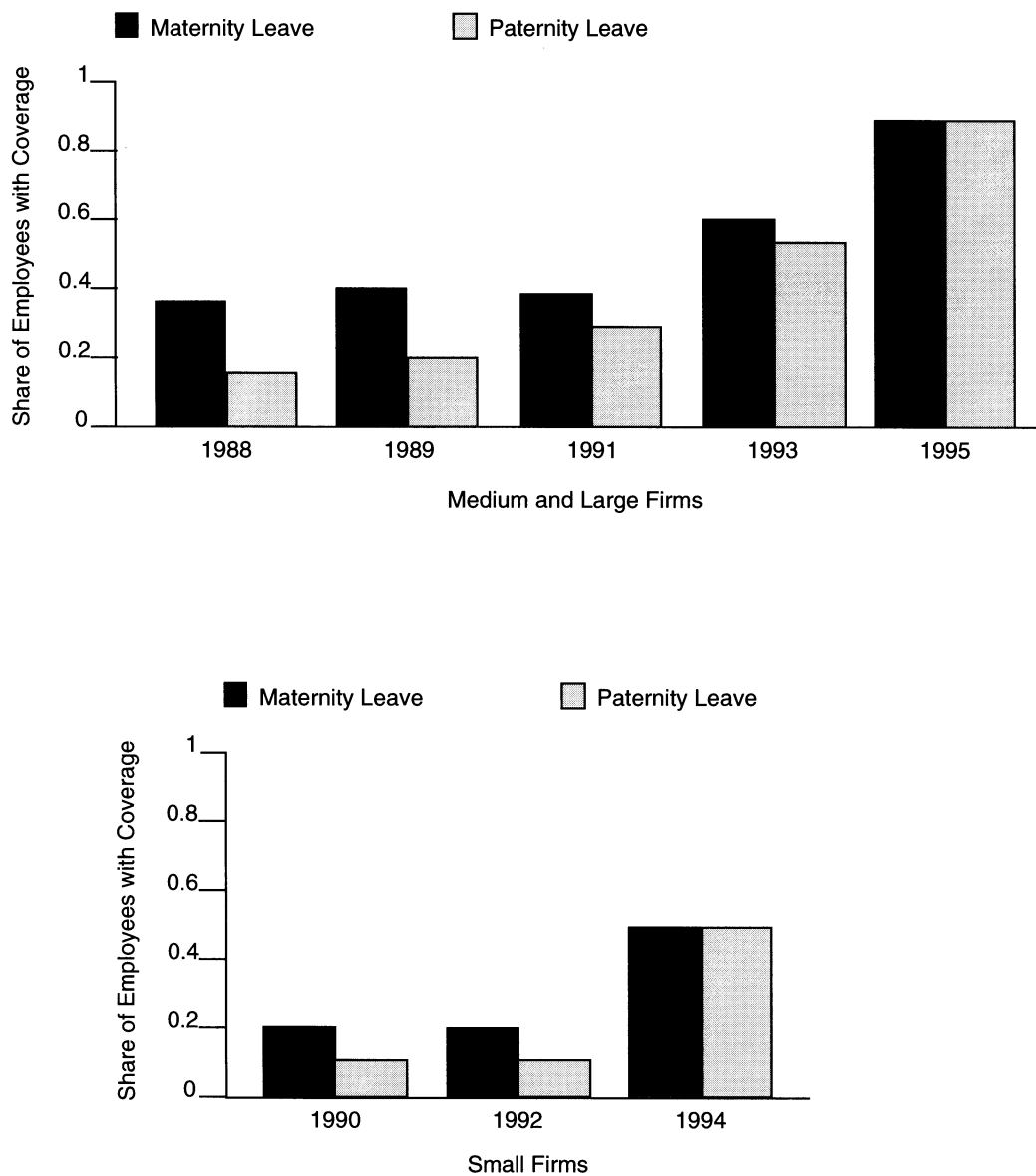


Figure 1. Increase in leave coverage by firm type in the employee benefits surveys, 1988–1995.

policies as a result of the FMLA, and two-thirds (67 percent) said they did.¹⁷ For the most part, the changes involved providing health insurance coverage during maternity or other types of leave (53 percent of firms making changes), providing job protection (54 percent), extending the period of leave (66 percent), or allowing other types of leave (77 percent).¹⁸

It appears that the FMLA had two main effects on coverage. First, these employer surveys suggest that some companies are now offering job-protected leaves for maternity and other family and medical reasons for the first time. Second, according to the Westat survey, many companies that previously offered leave have expanded their coverage (for example, by offering a longer leave, continuing health insurance during the leave, or allowing other types of leave).

QUESTION 2: DID UTILIZATION INCREASE AS A RESULT OF THE FMLA?

To address this question, I use data from the March 1992–1995 CPS.¹⁹ The 1992 and 1993 March CPS provide two years of data pre-FMLA (that is, prior to the implementation of the FMLA in August 1993); the 1994 and 1995 March CPS provide two years of data post-FMLA. Because the focus here is on women who have children and are in their childbearing years, I restrict the treatment group sample to women who have children and are aged 19 to 45; the control groups are men and childless women of the same age range. As noted earlier, I exclude part-time workers and the self-employed. The advantages of using the March CPS are well known; most important for the present analysis, it is a large, nationally representative sample, and it is the only month of the CPS that includes the necessary demographic and labor market variables (including the presence and age of children, current labor market activity, and employer characteristics such as firm size).²⁰

There are (at least) three reasons why we might observe a positive effect of the FMLA on leave-taking (which is defined here as the percentage of employees who have a job but are absent from work during the CPS survey week): (a) there might be new coverage by firms that did not previously offer coverage; (b) there might be expanded coverage by firms that already had coverage (if the FMLA was more generous than existing state law); and (c)

¹⁷ When the raw survey data are released to the public, it should be possible to calculate the number of employees whose employers changed policy as a result of the FMLA, the number of employees affected by various policy changes, and the like. It will also be possible to identify whether medium-sized firms were more likely than large firms to have changed their policies as a result of the FMLA.

¹⁸ Among the other types of leave offered, leave for fathers to care for newborn or sick children was the most frequently mentioned (69 percent of firms making changes). The survey also found that firms covered by the FMLA were more likely than uncovered firms to have leave policies and more likely to offer job protection as part of such policies.

¹⁹ The Family and Medical Leave Commission also sponsored a survey of employees, designed to assess the effect of the FMLA on leave-taking. This survey in fact did find that workers covered by the FMLA were more likely to take leave when they needed it [see McGonagle et al., 1995]. In the absence of data pre-FMLA, however, it is not possible to determine to what extent this difference is due to the FMLA.

²⁰ CPS revised some of its questions regarding labor force participation and absences from work in 1994, and there is some indication that the change in the survey may have resulted in changes in the reported levels of both employment and leave-taking. The potential for measurement bias provides another reason for using a difference-in-difference and difference-in-difference-in-difference methodology here.

there might be expanded usage in firms that already had coverage, due to an increased level of awareness about the available coverage. The new coverage effect should be stronger in states that had no state laws than in the others; the expanded coverage and usage effects should be observed in both. All three effects should be observed only for employees covered by the law, that is, those working full time and in firms with more than 50 employees, unless the passage of the law had indirect effects on those not covered (e.g., changing employer or employee attitudes about leave or employee expectations about benefit provision). Among firms covered by the law, the new coverage effect should be smaller in the very large firms (those with 500 or more employees) that were already most likely to offer coverage, but the expanded coverage and usage effects should be observed in all covered firms.²¹

One way to assess how leave-taking changed over time among the treatment and control groups, by firm size, is a simple difference-in-difference (D-D) and difference-in-difference-in-difference (D-D-D) analysis of the mean probability of an employee being on leave in each treatment group and in each control group in the states with and without laws, pre- and post-FMLA. These means are simply the coefficients from the following linear probability model:

$$LT_{ts} = \beta_1 exp*pre_{ts} + \beta_2 exp*post_{ts} + \beta_3 non*pre_{ts} + \beta_4 non*post_{ts} \quad (1)$$

where:

t indexes type of treatment or control group

s indexes firm size group

LT= leave-taking (1 if on leave during the CPS survey week, 0 otherwise);

*exp*pre*= year is 1992 or 1993 and state had no state law as of 1992

*exp*post*= year is 1994 or 1995, and state had no state law as of 1992

*non*pre*= year is 1992 or 1993, and state had a state law as of 1992

*non*post*= year is 1994 or 1995, and state had a state law as of 1992

The D-D estimate is $(\beta_2 - \beta_1) - (\beta_4 - \beta_3) = (\beta_2 - \beta_4) - (\beta_1 - \beta_3)$. The D-D-D estimate is the D-D estimate for a given treatment group and firm size group minus the D-D estimate for the control group from the same firm size group.²²

Table 3 shows the D-D and the D-D-D estimates of the effects of the FMLA on leave-taking. The D-D results for medium firms (100 to 499 employees) suggest that the FMLA did have a positive effect on leave-taking for both treatment groups (although the difference-in-difference is not significant for women with infants) but had no effect on leave-taking for the control groups; hence, the D-D-D estimates are positive (although not significant). The D-D-D estimates for the other firms look very different. In very small firms (1 to 24 employees), which are not covered by the FMLA, these estimates are all small and insignificant. In small firms (25 to 99

²¹ For evidence that firms with more than 500 employees were the most likely to have offered leave pre-FMLA, see Zigler and Frank [1988] and Hyde and Essex [1991]. Note that if the FMLA has had the intended effect of allowing employees to take leave rather than lose their jobs, it should also have a positive effect on employment (defined here as the percentage of the population who had a job during the CPS survey week, including both those at work and those absent from work), but I look at this in the following section because there may be offsetting negative employment effects as well.

²² There is no constant term in this model so that all the variation is apportioned to the year and state law group interaction terms.

Table 3. Difference-in-difference (D-D) and difference-in-difference-in-difference (D-D-D) estimate of the effect of the Family and Medical Leave Act (FMLA) on leave-taking among full-time employees.

Panel A. Treatment group 1: Women aged 19 to 45 in medium firms, with children under age 1.

	Pre-FMLA	Post-FMLA	Difference
No state law	3.20	16.08	12.88**
Had state law	5.17	11.76	6.59
Difference	-1.97	4.32	D-D: 6.29

Panel B. Treatment group 2: Women aged 19 to 45 in medium firms, with children under age 18.

	Pre-FMLA	Post-FMLA	Difference
No state law	1.07	3.59	2.52
Had state law	0.97	1.45	0.48
Difference	0.10	2.14**	D-D: 2.04**

Panel C. Control group 1: Women aged 19 to 45 in medium firms, with no children.

	Pre-FMLA	Post-FMLA	Difference
No state law	0.68	2.33	1.65**
Had state law	1.24	1.84	0.60
Difference	-0.56	0.49	D-D: 1.05

Panel D. Control group 2: Men aged 19 to 45 in medium firms.

	Pre-FMLA	Post-FMLA	Difference
No state law	0.61	1.72	1.11**
Had state law	0.72	1.74	1.02
Difference	-0.11	-0.02	D-D: 0.09

Panel E. Difference-in-difference-in-differences, medium firms (100 to 499 employees).

	Compared to men	Compared to childless women
1. Women with children under 1	(6.29-0.09) = 6.20	(6.29-1.05) = 5.24
2. Women with children under 18	(2.04-0.09) = 1.95	(2.04-1.05) = 0.99

Panel F. Difference-in-difference-in-differences, large firms (500 or more employees).

1. Women with children under 1	(-3.14-0.11) = -3.25	(-3.14+0.02) = -3.12
2. Women with children under 18	(-0.53-0.11) = -0.64	(-0.53+0.02) = -0.51

Panel G. Difference-in-difference-in-differences, small firms (25 to 99 employees).

1. Women with children under 1	(-9.58-0.37) = -9.95	(-9.58-1.80) = -11.38
2. Women with children under 18	(-0.54-0.37) = -0.91	(-0.54-1.80) = -2.34

Table 3. (continued)

Panel H. Difference-in-difference-in-differences, very small firms (1 to 24 employees).

	Compared to men	Compared to childless women
1. Women with children under 1	$(0.58+0.65) = 1.23$	$(0.58-0.15) = 0.43$
2. Women with children under 18	$(-0.99+0.65) = -0.34$	$(-0.99-0.15) = -1.14$

Notes: The Pre-FMLA estimates are from the March 1992 and 1993 Current Population Survey (CPS); post-FMLA estimates are from the March 1994 and 1995 CPS. Leave-taking for each group is defined as: (the number with a job but not at work during the survey week)/(the total number with a job that week). The D-D-D estimate in Panel E are from estimates for medium firms shown in panels A through D. The D-D-D estimates in Panels G and H are from D-D estimates (available upon request) for large, small, and very small firms, respectively. The difference-in-difference-in-differences is the difference-in-difference for the relevant treatment group minus the difference-in-difference for the relevant control group.

** Statistically significant at the 5-percent level.

employees) and large firms (500 or more), on the other hand, the D-D-D estimates tend to be negative, suggesting that states with prior state laws saw a bigger increase in leave-taking than those without state laws. The effect is particularly marked for the small firms. This latter result might reflect the fact that some of these firms were covered by state laws and that awareness and use of such coverage may have increased post-FMLA.

To control for other characteristics that might affect the probability of being on leave, I estimate the following probit model for leave-taking for each of the treatment groups and the control groups, again both separating employees by firm size and restricting the analysis to full-time workers:

$$P(LT) = F(\beta_0 + \beta_1 age + \beta_2 ed4 + \beta_3 ed3 + \beta_4 ed2 + \beta_5 married + \beta_6 prev + \beta_7 kidsu6 + \beta_8 kidsu18 + \beta_9 black + \beta_{10} hispanic + \beta_{11} exp + \beta_{12} y93 + \beta_{13} exp*y93 + \beta_{14} y94 + \beta_{15} exp*y94 + \beta_{16} y95 + \beta_{17} exp*y95) \quad (2)$$

where:

- $LT =$ 1 if on leave during the CPS survey week, 0 otherwise
- $age =$ age in years
- $ed4 =$ 1 if college degree or higher, 0 otherwise
- $ed3 =$ 1 if some college only, 0 otherwise
- $ed2 =$ 1 if high school only, 0 otherwise
- $married =$ 1 if currently married, 0 otherwise
- $prev =$ 1 if previously married, 0 otherwise
- $kidsu6 =$ number of children under age 6
- $kidsu18 =$ number of children under age 18
- $black =$ 1 if African American, 0 otherwise
- $hispanic =$ 1 if Hispanic origin: 0 otherwise
- $exp =$ 1 if state had no law as of 1992, 0 otherwise
- $y93 =$ 1 if year is 1993, 0 otherwise
- $exp*y93 =$ 1 if year is 1993 and state had no law as of 1992, 0 otherwise

Table 4. Marginal effects of the Family and Medical Leave Act (FMLA) on leave-taking, full-time employees, by firm size.

	Treatment Groups		Control Groups	
Medium firms (100 to 499 employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45
Post-FMLA, 1994	0.0664 (0.0486)	0.0116 (0.0121)	0.0006 (0.0075)	0.0137** (0.0062)
Post-FMLA, 1994 no state law	-0.0123 (0.0342)	0.0022 (0.0122)	0.0335** (0.0168)	0.0038 (0.0064)
Post-FMLA, 1995	-0.0031 (0.0326)	0.0058 (0.0118)	0.0028 (0.0078)	0.0121** (0.0061)
Post-FMLA, 1995 no state law	0.2335** (0.1157)	0.0372** (0.0189)	0.0340** (0.0170)	-0.0008 (0.0059)
Observations	377	5713	4538	12,521
Pseudo R ²	0.2137	0.0757	0.0361	0.0287
Large firms (500+ employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45
Post-FMLA, 1994	0.1841** (0.0715)	0.0418** (0.0098)	0.0148** (0.0069)	0.0142** (0.0044)
Post-FMLA, 1994 no state law	-0.0198 (0.0647)	-0.0079 (0.0069)	0.0052 (0.0075)	0.0010 (0.0042)
Post-FMLA, 1995	0.1582** (0.0686)	0.0392** (0.0094)	0.0214** (0.0071)	0.0178** (0.0044)
Post-FMLA, 1995 no state law	0.0091 (0.0678)	-0.0015 (0.0075)	0.0031 (0.0069)	0.0005 (0.0041)
Observations	1112	16,063	14,126	34,780
Pseudo R ²	0.0709	0.0648	0.0359	0.0276

y94= 1 if year is 1994, 0 otherwise
exp*y94= 1 if year is 1994 and state had no law as of 1992, 0 otherwise
y95= 1 if year is 1995, 0 otherwise; and
exp*y95= 1 if year is 1995 and state had no law as of 1992, 0 otherwise.

If the FMLA has had an effect on expanding utilization of leave regardless of prior state law, this would be seen in positive coefficients on the 1994 and 1995 year dummies, although clearly these year dummies will also pick up the effects of other national trends in those years. Positive coefficients on the interaction terms for 1994 and 1995 would indicate that the positive effect of the FMLA on leave-taking has

Table 4. (continued)

Small firms (25 to 99 employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45
Post-FMLA, 1994	0.0419 (0.0968)	0.0075 (0.0125)	0.0201 (0.0149)	0.0113* (0.0061)
Post-FMLA, 1994 no state law	-0.0382 (0.0850)	0.0067 (0.0146)	0.0132 (0.0169)	0.0051 (0.0066)
Post-FMLA, 1995	0.1018 (0.0957)	0.0298** (0.0141)	0.0098 (0.0140)	0.0216** (0.0068)
Post-FMLA, 1995 no state law	-0.0417 (0.0740)	-0.0038 (0.0112)	0.0176 (0.0193)	-0.0009 (0.0053)
Observations	271	4308	3882	12,339
Pseudo R^2	0.1161	0.0409	0.0575	0.0465
Very small firms (1 to 24 employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45
Post-FMLA, 1994	0.0148 (0.0805)	0.0168* (0.0088)	0.0114 (0.0084)	0.0223** (0.0066)
Post-FMLA, 1994 no state law	0.1239 (0.1151)	0.0045 (0.0094)	-0.0034 (0.0081)	-0.0031 (0.0051)
Post-FMLA, 1995	0.0576 (0.0915)	0.0221** (0.0094)	0.0098 (0.0088)	0.0323** (0.0073)
Post-FMLA, 1995 no state law	0.0446 (0.1198)	0.0061 (0.0098)	0.0082 (0.0105)	-0.0065 (0.0046)
Observations	293	6076	5560	19,046
Pseudo R^2	0.0965	0.0853	0.0383	0.0427

Notes: Marginal effects (with standard errors in parentheses) are from probit models which also include controls for age, college plus, some college, high school only, married, previously married, number of children under 6, number of children under 18, African American, Hispanic, as well as dummy variables for whether the state had a law in 1992, whether the year is 1993, and the interaction of state law and 1993.

* Statistically significant at the 10-percent level.

** Statistically significant at the 5-percent level.

been greater in the states without prior state laws than in the states that had state laws. Negative coefficients on the interaction terms would suggest the opposite, that is, that the increase in leave-taking has been bigger in states with state laws.²³

²³ This model has the advantage of being readily interpretable, although it is not directly comparable with the D-D-D models shown in Table 3. A D-D-D regression model could be estimated, but it would be computationally quite cumbersome as it would have to allow for separate coefficients by group (because the returns to demographic and human capital characteristics are not identical across groups).

Table 4 shows the marginal effects of the FMLA on leave-taking among employees of large, medium, small, and very small firms, from probit models run separately for the two treatment groups (women with children under 18 and women with children under 1) and the two control groups (men and childless women). The results for medium firms confirm that there indeed was an effect of the FMLA in covered firms, as there was an increase in leave-taking for both treatment groups in 1995 in states that did not have a law prior to the FMLA, with a particularly marked increase of 23 percent for women with children under age 1. In contrast, there was no such increase for men in medium firms and only a small increase for childless women. This result is consistent with the expectation that there would be a new coverage effect concentrated among women with infants, in medium firms, in states with no prior law. In large firms, on the other hand, there were substantial increases in leave-taking for women with infants across all states in 1994 and 1995, small increases for women with children overall, and smaller increases for men and childless women. This is consistent with an expanded coverage and usage effect in firms that were likely to already have had coverage pre-FMLA; there may also have been a secular increase in leave-taking unrelated to the law. Small firms display a different pattern, with some evidence of an increase in leave-taking in 1995 and, as in the D-D-D estimates, some negative (but insignificant) effects of being in a state without a state law, suggesting that state laws continue to be more important than the FMLA for these firms. Finally, in very small firms, as expected there are only small increases in leave-taking.

The difference in the impact of the FMLA on leave-taking among women with children and women with infants in medium firms and large firms, as compared to men and childless women in the same-sized firms, is particularly striking. The FMLA apparently has little effect in the large firms; this makes sense because these firms are more likely to provide leave regardless of state or federal law. Medium-size firms, on the other hand, are apparently more sensitive to legislation in this area. As a result, the FMLA has made a big difference to the medium size firms in states where there was no prior law in effect.

QUESTION 3: DID THE FMLA HAVE AN EFFECT ON WOMEN'S EMPLOYMENT?

The likely direction of employment effects is not clear a priori. If women's utilization of leave has increased subsequent to the FMLA and if women are more likely to take leave and return to their jobs as opposed to leaving their jobs altogether, then we would expect to see a positive effect on employment. However, if employers pass along the costs of the mandated benefit to those potentially benefiting from it (in this case, women of childbearing age), then we might observe a negative employment effect (unless wages are fully flexible). Further, if these two effects offset each other, we might observe no net effect at all.

To address this question, I use the March 1992–1995 CPS. States with no prior laws constitute the experimental group; states with prior laws are the nonexperimental group. As was true earlier, women who are aged 19 to 45 and have children, or have infants, are considered treatment groups. In addition, childless women aged 19 to 45 now constitute another treatment group, because they are potential beneficiaries of the law and might bear some of the costs of the law. The control groups are men aged 19 to 45 and women aged 46 to 60; both of these groups are presumably less likely to take advantage of the law than are women in their childbearing years. The effects are not analyzed separately by firm size, as this is not observed for those not employed. This is an important limitation, as the effects of the FMLA should vary by firm size.

In the D-D analysis (not shown), there is an upward trend in the employment

Table 5. Marginal effects of the Family and Medical Leave Act (FMLA) on employment.

	Treatment Groups			Control Groups	
	Women aged 19 to 45, kid < 1	Women aged 19 to 45, kid < 18	Women aged 19 to 45, no kids	All men aged 19 to 45	All women aged 46 to 60
Post-FMLA, 1994	-0.0088 (0.0325)	0.0030 (0.0101)	-0.0208** (0.0103)	-0.0045 (0.0061)	0.0081 (0.0126)
Post-FMLA, 1994 no state law	0.0373 (0.0380)	0.0145 (0.0116)	0.0142 (0.0116)	0.0154** (0.0070)	0.0050 (0.0146)
Post-FMLA, 1995	0.0761** (0.0325)	0.0215** (0.0101)	-0.0158 (0.0104)	0.0167** (0.0060)	0.0123 (0.0125)
Post-FMLA, 1995 no state law	-0.229 (0.0380)	0.0070 (0.0118)	0.0124 (0.0118)	0.0057 (0.0072)	0.0027 (0.0145)
Observations	7637	75,189	48,622	111,757	47,045
Pseudo R^2	0.1017	0.1038	0.0778	0.0764	0.0861

Notes: Marginal effects (with standard errors in parentheses) are from probit models that also include controls for age, college plus, some college, high school only, married, previously married, number of children under 6, number of children under 18, African American, Hispanic, and other family income as well as dummy variables for whether the state had a law in 1992, whether the year is 1993, and the interaction of state law and 1993. Employment is defined as being employed as of the survey week.

** Statistically significant at the 5-percent level.

of women with infants in both sets of states, but the difference in the employment growth of women with infants between states that had a law and those that did not is not significant (the D-D-D estimates are insignificant as well). In fact, the only significant D-D is the larger employment growth for women overall in states with no prior law.

To control for other characteristics that might affect employment, I use a probit model, as follows:

$$\begin{aligned}
 P(E) = & F(\beta_0 + \beta_1 age + \beta_2 ed4 + \beta_3 ed3 + \beta_4 ed2 + \beta_5 married + \beta_6 prev \\
 & + \beta_7 kidsu6 + \beta_8 kidsu18 + \beta_9 black + \beta_{10} hispanic + \beta_{11} othinc \\
 & + \beta_{12} exp + \beta_{13} y93 + \beta_{14} exp*y93 + \beta_{15} y94 + \beta_{16} exp*y94 \\
 & + \beta_{17} y95 + \beta_{18} exp*y95)
 \end{aligned} \quad (3)$$

where E equals employment (1 if had a job during the survey week, whether at work or not; 0 otherwise); $othinc$ equals other family income; and all the other variables are defined as earlier.

The marginal effects from the employment probits are shown in Table 5. The first noteworthy result is that there is little evidence of any employment effect for any of the treatment groups in the states without prior laws as opposed to those with prior laws; the only significant interaction between year and state without a law is the one

for all men in 1994, and that coefficient is quite small. This result suggests that any negative employment effects for women overall were either negligible or were offset by positive effects.

The second result of interest is the 7.6 percent increase in the employment of women with children under the age of 1 in 1995, notably larger than the increase in employment for any other group in that year. Because this increase is seen in both sets of states, it may reflect the employment effects of expanded coverage and usage, or the effects of other national trends in employment in those years. The fact that the increase is not larger in the states without prior laws suggests that if there were an effect of the FMLA on new coverage, the positive employment effects associated with this increase in coverage may have been offset by negative employment effects as employers adjusted to the costs of coverage.

QUESTION 4: DID THE FMLA HAVE AN EFFECT ON WOMEN'S WAGES?

As with employment, the effect of the FMLA on the wages of affected groups is unclear a priori. One hypothesis is that wages of women of childbearing age will go down, as employers pass along the costs of the mandated benefit to those potentially benefiting from it [see Burtless, 1995; Summers, 1989]. To the extent that women valued the benefit, they might be willing to accept lower wages in return for the benefit.²⁴ If this were the case, we should see a fall in women's wages relative to men's in the states that did not have a prior law as compared to those that did. A second hypothesis is that wages of women using the FMLA, in particular women with infants and more generally women with children, will go up, as women use the FMLA to retain their jobs, thus receiving a wage premium associated with seniority and a good job match.²⁵ In this case, we might see a rise in the wages of women with infants and women with children overall, although it might be too soon to see this effect if the wage gains increase over time. Again, if the first of these effects offsets the other, the net effect observed might be zero.

Wages from the CPS are estimated by dividing last year's wage and salary income by the product of weeks worked last year and usual hours worked per week. Thus, 1994 wages actually reflect earnings in calendar year 1993, when the FMLA was just coming into effect. For this reason, I place greater emphasis on the 1995 results in the regression models that follow.

In the D-D and D-D-D estimates (not shown), there is an increase in wages for women with infants working in small or medium firms which is smaller in the states without prior laws than in the other states, but the difference is significant only in the small firms.

Because women who work post-FMLA may differ in characteristics (such as education) that are correlated with their wages, it is particularly important to examine these wage effects controlling for other characteristics. Using Ordinary Least Squares (OLS), I estimate a basic human capital earnings function, as follows:

$$\begin{aligned}
 W = & \beta_0 + \beta_1 age + \beta_2 ed4 + \beta_3 ed3 + \beta_4 ed2 + \beta_5 married + \beta_6 prev \\
 & + \beta_7 kidsu18 + \beta_8 black + \beta_9 hispanic \\
 & + \beta_{10} exp + \beta_{11} y93 + \beta_{12} exp*y93 + \beta_{13} y94 + \beta_{14} exp*y94 \\
 & + \beta_{15} y95 + \beta_{16} exp*y95
 \end{aligned} \tag{4}$$

²⁴ For a recent discussion of compensating differentials and mandated benefits, see Anderson and Meyer [1996].

²⁵ See Waldfogel [1997, 1998b], who uses panel data to examine these effects.

where W equals the log of hourly wage, in 1995 dollars. (All the other variables are defined as in the earlier models.) This model, like the previous regression models, includes a fixed-state effect (for experimental vs. nonexperimental states), which is particularly important here because the states with prior state laws tend to have higher mean wages than those without state laws.

The results of the log wage regressions, shown in Table 6, present a mixed picture as to the two hypotheses outlined earlier. If there were substantial cost-shifting (with no offsetting effects), we should see it for women in medium firms in states with no

Table 6. The effect of the Family and Medical Leave Act (FMLA) on the log of hourly wages, full-time employees, by firm size.

	Treatment Groups			Control Groups	
	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45	All women aged 46 to 60
Medium firms (100 to 499 employees)					
Post-FMLA, 1994	-0.0923 (0.1307)	0.0024 (0.0336)	-0.0156 (0.0352)	-0.0100 (0.0219)	0.0751 (0.0605)
Post-FMLA, 1994 no state law	0.0524 (0.1528)	0.0150 (0.0389)	-0.0283 (0.0423)	-0.0205 (0.0259)	-0.0810 (0.0718)
Post-FMLA, 1995	-0.0464 (0.1232)	0.0004 (0.0328)	-0.1015** (0.0365)	-0.0174 (0.0221)	0.0458 (0.0604)
Post-FMLA, 1995 no state law	-0.0167 (0.1456)	0.0087 (0.0384)	0.1021** (0.0436)	-0.0211 (0.0261)	-0.0241 (0.0717)
Observations	373	5632	4474	12,386	2255
Adj. R^2	0.3321	0.2653	0.2261	0.3101	0.2966
Large firms (500+ employees)					
Post-FMLA, 1994	-0.0588 (0.0775)	0.0012 (0.0212)	-0.0268 (0.0209)	-0.0497** (0.0139)	-0.0040 (0.0383)
Post-FMLA, 1994 no state law	0.1151 (0.0890)	0.0094 (0.0242)	-0.0081 (0.0246)	0.0180 (0.0161)	-0.0186 (0.0454)
Post-FMLA, 1995	-0.1840 (0.0776)	-0.0192 (0.0209)	-0.0830** (0.0206)	-0.0435** (0.0137)	0.0173 (0.0376)
Post-FMLA, 1995 no state law	0.1671* (0.0886)	0.0049 (0.0238)	0.0603** (0.0242)	0.0209 (0.0158)	-0.0497 (0.0446)
Observations	1098	15,846	13,915	34,472	6738
Adj. R^2	0.3974	0.2560	0.2692	0.3203	0.2215

Table 6. (continued)

Small firms (25 to 99 employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45	All women aged 46 to 60
Post-FMLA, 1994	0.3345** (0.1613)	0.0852** (0.0403)	0.0085 (0.0391)	-0.0122 (0.0228)	-0.0063 (0.0777)
Post-FMLA, 1994 no state law	-0.3684 (0.1830)	-0.0765 (0.0462)	-0.0355 (0.0465)	0.0199 (0.0271)	0.0537 (0.0929)
Post-FMLA, 1995	0.2948** (0.1471)	0.1057** (0.0402)	-0.0274 (0.0407)	-0.0197 (0.0226)	-0.1212 (0.0779)
Post-FMLA, 1995 no state law	-0.2588 (0.1712)	-0.1028* (0.0463)	0.0295 (0.0477)	0.0109 (0.0270)	0.1005 (0.0931)
Observations	266	4208	3876	30,479	1702
Adj. R ²	0.3586	0.2310	0.2280	0.2633	0.1874
Very small firms (1 to 24 employees)	Women aged 19 to 45, kids < 1	Women aged 19 to 45, kids < 18	Women aged 19 to 45, no kids	All men aged 19 to 45	All women aged 46 to 60
Post-FMLA, 1994	0.2783** (0.1321)	0.0594 (0.0378)	0.0360 (0.0348)	-0.0450** (0.0207)	0.0180 (0.0906)
Post-FMLA, 1994 no state law	-0.2427 (0.1596)	-0.0283 (0.0434)	-0.0184 (0.0416)	0.0659** (0.0244)	0.0389 (0.1031)
Post-FMLA, 1995	0.0731 (0.1438)	-0.0119 (0.0393)	-0.0179 (0.0364)	-0.0691** (0.0212)	-0.0202 (0.0920)
Post-FMLA, 1995 no state law	-0.0441 (0.1738)	0.0112 (0.0450)	-0.0114 (0.0431)	0.0762** (0.0249)	0.0553 (0.1046)
Observations	349	5685	5436	18,309	2339
Adj. R ²	0.2847	0.1421	0.1702	0.2433	0.1140

Notes: Standard errors are in parentheses. The dependent variable is the log of hourly wages. All wages are in 1995 dollars. Wages are calculated by dividing last year's reported earnings by last year's hours worked (weeks worked last year times usual hours worked per week). Extreme values (wages below 50 percent of the minimum wage or above \$400/hour) are excluded. Models are estimated using OLS and also include an intercept, age, college degree or more, some college only, high school only, currently married, previously married, number of children under 18, African American, Hispanic, as well as dummy variables for whether the state had a law as of 1992, whether the year is 1993, and the interaction of state law and 1993.

* Statistically, significant at the 10-percent level.

** Statistically significant at the 5-percent level.

prior law and for women in large firms (because these are the two sectors where leave-taking increased the most) but we do not. In terms of positive wage effects, whether due to the retention of women in their jobs or the selection of more qualified women into employment, again, looking at the groups with the biggest increases in leave-taking, none of the wage effects are positive (although arguably the zero effect for women with infants in medium firms with no state law could be seen as the result of offsetting wage effects). In the large firms, where leave-taking increased in both sets of states in 1995, we find positive wage effects only in states with no prior laws, which suggests that the positive wage effects may outweigh the negative effects in large firms.

Oddly, the only groups where negative wage effects of the FMLA emerge are the small and very small firms, where the wages of women with infants rise sharply in both 1994 and 1995, but with large negative effects offsetting these increases in the states with no prior laws. This raises the possibility that at least some small and very small firms granted more leave after the FMLA even though they were not covered by it and that, in states where this was a new mandate, firms reduced wages accordingly; however, we see no evidence of this in the leave-taking data. A more likely interpretation is that employees in small and very small firms in states with prior state laws were more likely to take job-protected leave and then return to work after the passage of the FMLA because of increased publicity and awareness of their rights under existing state law. The higher wages for women with infants and women overall in those states, then, could possibly be the result of their being in better job matches and having higher job tenure as a result of having taken leave and having kept their jobs.²⁶

CONCLUSIONS

This article provides evidence on the impact of the FMLA on coverage, leave-taking, employment, and earnings. The results cast doubt on two commonly held assumptions about the FMLA.

First, it is commonly thought that the FMLA probably has had little effect on coverage and leave-taking because it excludes so many workers and firms, because those it covers were probably covered already, and because the leave is unpaid. Although it is true that the impact has been limited, there has nevertheless been a notable increase in coverage and an even bigger increase in leave-taking. The increase in leave-taking appears to have two sources. One source is new usage among those newly covered. The second source is greater usage among those already covered, which occurs because the FMLA has expanded coverage and because it may also have affected employer and employee attitudes about leave-taking.²⁷

Second, it is often assumed that the FMLA might have a negative effect on the employment and earnings of affected groups because it is a mandated benefit that imposes costs on employers. In the case of the FMLA, however,

²⁶ The latter explanation could be tested more directly using panel data such as the NLSY; this is a promising direction for further research.

²⁷ Two new studies [Klerman and Leibowitz, 1998; Ross, 1998] provide additional evidence of greater usage among those already covered; both studies found that maternity leave lengths have increased among women newly covered by the FMLA.

the cost of the mandate is fairly low, and there are also likely to be positive employment and earnings effects associated with the mandate. This article finds that these effects to a large extent offset each other, resulting in slightly positive net employment effects and essentially zero net wage effects (at least in the short run).²⁸

What do these results suggest about possible extensions of the FMLA? One frequent suggestion is to change the law to make some provision for paid leave.²⁹ Although this article provides no direct evidence on paid leave, the results suggest that the FMLA has had no (net) negative employment or wage effects thus far even though it has imposed some costs on employers. Providing full or partial wage replacement for leave would entail much higher costs for those firms that do not currently provide paid leave.³⁰ Mandating paid leave could therefore have negative effects on employment or wages, but again the net effect might be zero given the likely offsetting positive employment and wage effects.

The other frequently suggested reform of the FMLA is to extend coverage to the roughly 50 percent of American workers in the private sector who are not currently covered by the law (those who work in smaller firms and those who work part time).³¹ The results in this article suggest that the net negative employment and wage effects of extending coverage to the remainder of the workforce are likely to be minimal, while the benefits in terms of new coverage and expanded usage could be substantial. Moreover, family leave legislation may provide other benefits, such as improved health outcomes for mothers and children. The effect of the FMLA on these outcomes is a promising area for further research.

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²⁸ In recent research on state maternity leave laws, Kallman [1996] also finds positive employment effects; however, she finds negative wage effects associated with the state laws. Ruhm [1997], in a review of the research on family leave, concludes that the FMLA would likely have small positive effects on affected groups' employment and wages.

²⁹ The Commission on Family and Medical Leave [1996] found that two-thirds of leave-takers receive at least partial payment during their leave, with nearly half receiving full payment. Those who do not have paid leave, however, are more likely to be from lower-income families and are thus least able to manage without pay. For example, 12 percent of women leave-takers turn to public assistance to replace the income they lose while on unpaid leave.

³⁰ Another possible reform would be to adopt a social insurance model, with leave funded by employer and employee contributions. This is the model used by most other countries, but currently only five U.S. states have a similar program (temporary disability insurance); see Hartmann et al. [1995] and Waldfogel [1998b].

³¹ In principle, it might be possible to assess the likely impact of these extensions by comparing states with laws that cover small firms and part-time workers with those that do not, but the small number of states with such laws would make such a comparison speculative at best.

APPENDIX

Table A.1. States with laws providing job-protected maternity leaves, in effect at least one year prior to the Family and Medical Leave Act (FMLA).

State	Month/Year	Firm size covered
California	1/80	15 ⁺ until 1/92; 1 ⁺ after
Connecticut	1/73	3 ⁺
Maine	4/88	25 ⁺
Massachusetts	10/72	6 ⁺
Minnesota	7/87	21 ⁺
New Jersey	4/90	75 ⁺ until 5/93; 50 ⁺ after
Oregon	1/88	25 ⁺
Rhode Island	7/87	50 ⁺
Tennessee	1/88	100 ⁺
Washington State	10/73	8 ⁺
Washington, DC	4/91	50 ⁺
Wisconsin	4/88	50 ⁺

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