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Does the Family and Medical Leave Act (FMLA) Increase Fertility Behavior?

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Abstract The Family and Medical Leave Act (FMLA), implemented in August 1993, grants job-protected leave to any employee satisfying the eligibility criteria. One of the provisions of the FMLA is to allow women to stay at home for a maximum period of 12 weeks to give care to the newborn. The effect of this legislation on the fertility response of eligible women has received little attention by researchers. This study analyzes whether the FMLA has influenced birth outcomes in the U.S. Specifically, I evaluate the effect of the FMLA by comparing the changes in the birth hazard profiles of women who became eligible for FMLA benefits such as maternity leave, to the changes in the control group who were not eligible for such leave. Using a discrete-time hazard model, results from the difference-in-differences estimation indicate that eligible women increase the probability of having a first and second birth by about 1.5 and 0.6 % per annum, respectively. Compared to other women, eligible women are giving birth to the first child a year earlier and about 8.5 months earlier for the second child.

Keywords Family and medical leave act · FMLA · Fertility · Births · Hazard models · Maternity leave · Difference-in-differences

JEL classification I18 · J00 · J13 · J18

Introduction

Since the 1990s, family-friendly policies such as maternity leave have increased throughout the United States. Universal maternity leave has become an important topic of discussion regarding its contribution to various aspects of work and family life, in particular, the effect on decisions related to childbearing (Noble 1993; Ife and Jen Kelly 2008; Brown 2009; Folbre 2010; Saltzman 2010).

For most countries, especially those in Europe such as Austria and Sweden, pronatalist policies—those designed to encourage fertility behavior—are quite common

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where parental leave is sufficiently generous allowing mothers to have two or possibly three separate gestations within the same leave period. In addition, these countries have provisions ensuring that employees are compensated while on leave. This is in sharp contrast to the United States (U.S.) where leave policies only became universal through the passage of the Family and Medical Leave Act (FMLA) in 1993.[1] The FMLA provides a shorter leave period amounting to 12 weeks unpaid leave, which is only one-third the average amount in most European and OECD countries.[2] Further, new mothers in Europe get on average 14 to 16 weeks of paid maternity leave (Kamerman and Gatenio 2002). Prior to the FMLA, only 12 states in the U.S. (plus the District of Columbia) possessed laws requiring job-protected maternity leave but with wide variation (from 4 to 18 weeks) in the length of time allowed.[3] An advantage of the FMLA over the state-specific leave laws is that it is universal and so better able to influence childbearing preferences at the national level.

Outside of the U.S., such as Canada (Milligan 2002) and Europe (Gauthier 2007), incentive-driven policies have been shown to have an impact on fertility rates. In line with this research are those studies that examine the impact on family fertility of work-related policies such as maternity leave. Nearly all of the studies use micro-level data, producing mixed results: some showing that work-related incentives have a positive impact in raising fertility (Hyatt and Milne 1991; Buttner and Lutz 1990; Hoem 1993; and Lalive and Zweimüller 2009) while others have found no evidence of a policy impact (Gauthier and Hatzius 1997 and Hoem et al. 2001).

Despite the evidence of some positive impact of family leave polices on family formation, the fertility effect of maternity leave in the U.S. under the auspices of the FMLA has been largely ignored within the empirical literature.[4] This neglect is perplexing. Consequently, this is the first analysis to study the fertility effects of a federal government-mandated parental leave policy in the U.S. In this paper, I estimate the impact of the Family and Medical Leave Act (FMLA) on fertility outcomes. Specifically, I examine whether eligible women under the FMLA criteria are more likely to have a first and second birth compared to other women. I find evidence that implementation of the FMLA is associated with an increase in fertility outcomes among eligible women. Using panel data from the 1979 cohort of the National Longitudinal Survey of Youth (NLSY79) for the period 1989 to 2010, I find that the introduction of the FMLA has significantly increased the probability of eligible women having a first and second child. The results are robust to a variety of specifications. I also find that the impact of the legislation is evident amongst eligible whites and those with college-level experiences. These results are corroborated with the findings that eligible women are giving birth faster as a result of the policy.

[1] Although it is a federal policy, knowledge of the FMLA is still not as widespread since only 58.2% of employees at covered worksites have any knowledge of FMLA (Department of Labor 2007).

[2] The United Nations watchdog on labor issues, the International Labor Organization (ILO) prescribes a minimum of 14 weeks maternity leave along with some cash benefit (International Labor Organization 2000).

[3] The states are California, Connecticut, District of Columbia, Maine, Massachusetts, Minnesota, New Jersey, Oregon, Rhode Island, Tennessee, Washington, Wisconsin and Vermont. Information obtained from Waldfogel (1999).

[4] Averett and Whittington (2001) investigate the impact of maternity leave on influencing births during the period 1985 to 1992, which was a period before the FMLA was introduced.

I organize the rest of the paper as follows: Section 2 provides background information on the FMLA. Section 3 outlines the previous empirical literature. Section 4 presents the empirical strategy while Section 5 presents the data. The results are presented in Section 6, while Section 8 discusses and concludes.

Background on the Family and Medical Leave Act (FMLA) of 1993

On February 5, 1993, President Bill Clinton signed his first piece of legislation by enacting the Family and Medical Leave Act (FMLA) of 1993, a federal law guaranteeing job-protected maternity leave along with other provisions. Prior to the FMLA, job-protected maternity leave—the right to take time off from work to give birth, care for the newborn and resume employment at the pre-birth job—constituted part of the broad compensation package decided between employees and employers. Presumably, employers need to attract and keep good workers so they rarely just pay minimum wage and minimum benefits. Mandates (minimum wage; maternity benefits) thus only matter to the extent they exceed the privately negotiated mix of compensation.[5] In large part, employer and state policies towards maternity leave were aimed at providing greater flexibility for women in the workplace and to strike some balance between work and parental duties.

Implemented on August 5, 1993, the FMLA provided benefits only to eligible employees. Under the FMLA, an eligible employee is one who satisfied the following criteria:

(1) has been employed by a covered employer;[6]
(2) has worked for the employer for at least 12 months;
(3) has worked at least 1,250 h over the previous 12 months; and
(4) is employed at a worksite in the United States or in any territory or possession of the United States where 50 or more employees are employed by the employer within 75 miles of that worksite.

Eligible employees are entitled to take up to 12 weeks of unpaid, job-protected leave each year for any of the following reasons: (a) to give birth and care for the newborn; (b) to adopt a child or to carry out foster care; and (c) to care for self, a spouse, a child or parent with a serious health condition.[7] The FMLA applies to all public agencies including those of the state, local and federal governments as well as schools. Employees in the private sector are also covered.

Previous Literature

Over the past decade, empirical findings have improved our knowledge about the effects of maternity leave as well as other family policies. There is a growing literature

[5] Less than 25% of the contiguous states of the U.S. had any state legislation establishing job-protected maternity leave.
[6] See U.S. Department of Labor website (http://www.dol.gov/whd/fmla/CONSADReport.pdf, accessed May 14, 2012) for a detailed explanation of what constitutes a covered employer.
[7] Under the FMLA, if two spouses are employed by the same employer, then the combined leave allowed is 12 weeks.

providing evidence on the impact of maternity leave policies on the labor market outcomes of women, more so in the U.S. An important aspect of maternity leave policies is a reduction in some of the wage gaps between mothers and non-mothers (Waldfogel 1998; Klerman and Leibowitz 1999). With respect to the impact of the FMLA on job-continuity, Baum (2003) finds a relatively small effect of this federal policy, primarily because prior leave policies were already in place in some firms. Some studies such as Waldfogel (1999), find that a higher proportion of women with very young children (less than one year) have been employed since the FMLA. In the same study, the author, investigating the impact on wages, finds that the impact of leave policies on wages is even less clear. On the aspect of the role of parental leave on leave-taking, there has been limited research. The more recent research in this area (example: Waldfogel 1999; Klerman and Leibowitz 1998) finds an increase in the amount of leave taken as a result of these leave policies.

There is a strand of literature that assesses the role of parental leave policies on the health outcomes of new mothers and their children. Using the U.S. data, Berger et al. (2005) find that women taking less than the maximum allowable leave period under the FMLA have children with worse health outcomes, attributable in part to the lower levels of breastfeeding and immunizations.

Parental and family leave policies in other countries such as Canada and Europe are paid and it is not surprising that these policies have a bigger impact than unpaid policies. For the impacts on fertility see Hyatt and Milne (1991), Buttner and Lutz (1990), Hoem (1993), Gauthier and Hatzius (1997), Hoem et al. (2001) and Lalive and Zweimüller (2009); female labor supply (Winegarden and Bracy 1995; Ruhm 1998; Baker and Milligan 2008a, 2008b); and child health (Winegarden and Bracy 1995; Ruhm 2000; Tanaka 2005).

Previous empirical research provides evidence that the introduction of the FMLA has led to more coverage for eligible recipients who have increased the amount of leave taken (Ross 1998; Waldfogel 1999; Han and Waldfogel 2003; Han et al. 2007). At the same time, it appears that the FMLA has not caused any significant changes in employment or wages. Berger and Waldfogel (2004) showed that FMLA-eligible women were more likely to access parental leave and then return to work after giving birth—a finding that may be attributable to the FMLA.

To date and to my knowledge, only a couple of published multivariate study has analyzed the impact of maternity leave on fertility outcomes in the U.S. Only one has investigated the impact of the FMLA on births among other outcomes (Rossin 2011), finding that the FMLA had a positive and statistically significant impact on first-births. The study, which was done by Averett and Whittington (2001) examines whether employer-provided maternity leave affects fertility decisions of women. Specifically, they investigated the impact of maternity leave on the probability of observing a birth using a discrete-time hazard model. They find an increase in the likelihood of a woman giving birth as a result of maternity leave.

If the introduction of the FMLA has created an incentive for more women to access leave (including maternity leave) then the FMLA may have impacted fertility behavior. This study expands on previous literature by investigating the causal effects of a universal policy (federal government legislation) on fertility behavior of women since the introduction of the FMLA. The research contributes to the existing literature in several distinct ways.

This is the first analysis to study the fertility effects of a federal government-mandated parental leave policy in the U.S. using longitudinal data. While the study by Averett and Whittington (2001) examines the fertility effects of employer-based maternity leave in the U.S., this was prior to the implementation of the FMLA. One limitation of their study is that only a few states had laws providing job-protected maternity leave prior to the FMLA. Rossin's (2011) study uses cross-sectional data from which she was only able to arrive at 'crude' measures of key variables related to the FMLA. Therefore, the present study makes it possible to undertake a perhaps more precise investigation of the universal fertility impact of a federal law. Consequently, the evidence that this paper provides on the impact of maternity leave-related policies on fertility outcomes may have broader policy implications. Second, I am able to exploit the variation in exposure to FMLA benefits. There is also variation in covered employers within states because not all firms are covered under FMLA. Further, two individuals in the same covered firm may differ in eligibility status if one of them does not meet the criteria explained in Section 2. This allows for the use of a difference-in-differences (DID) strategy.

Third, I use a dataset, specifically the NLSY79 which allows for the tracking of the behavior of a group of individuals over time and whose age group makes them most likely to be the first to be affected by this policy.[8] The data sample, which spans the survey period 1989 to 2010, provides important information over a sufficiently long time frame allowing for an assessment of the FMLA. Ruhm (1997) expresses a similar sentiment by informing that over time, usage of the FMLA will likely increase.

Empirical Strategy

Since identification comes from an exogenous policy rule (eligibility criteria) based on length of current employment and company size, it is necessary to identify the policy impact by comparing the likelihood of a birth outcome between individuals eligible for FMLA benefits and those who are ineligible. [9] To do this, I employ a difference-in-differences estimator exploiting the variation in exposure to FMLA benefits by identifying changes in fertility behavior between two groups. One group receives treatment and another group does not. The natural control group comprises women who were not eligible for FMLA, while the treatment group comprises women who satisfied the criteria for FMLA benefits and so became eligible.

To evaluate the impact of the policy, I compare the changes in risk or probability of giving birth in the treatment and control group using the following discrete-time hazard model:

$$\mathrm{logit}(\mathrm{h}_{it}) = \alpha(\mathrm{t}) + \beta_1 \mathrm{FMLA}_t + \beta_2 \mathrm{Eligible}_i + \beta_3 (\mathrm{FMLA}_t * \mathrm{Eligible}_i) + \beta_4 \mathbf{X}_{it} + e_{it} \quad (1)$$

[8] Waldfogel (1999) noted that the NLSY data as being a potentially good source of information in investigating the FMLA's impact on leave coverage. At the time of her research, the span of the data was too short to carry out any meaningful investigation.

[9] Ineligible women are those who do not satisfy all the eligibility criteria and who have worked some positive number of hours in each of the years during the survey period 1989–2010. Therefore, women who never worked in any year during this time period are excluded.

where h_{it} is the hazard rate for individual i during interval t. The hazard rate is the probability that an event occurs at a certain time to a particular individual given that the individual is at risk to the event at that time. If the event is a first birth, then the hazard rate is the probability of a woman giving a first birth within the survey period among those women who are yet to have a first birth.

The right-hand side of the equation consists of an indicator variable $FMLA_t$ which equals one for the years 1993 to 2010 and zero otherwise; $Eligible_i$ is a dichotomous variable which equals one if a woman satisfies all the eligibility criteria (defined in Section 2) in the first year of the FMLA and zero otherwise; while $FMLA_t$ *$Eligible_i$ is the interaction term. The vector X contains a set of covariates which may be fixed or time-varying and controls for age, race and ethnicity, marital status, education, family income, whether the woman has an urban residence, job tenure and the number of children the woman has had previously. The βs represent the corresponding parameters to be estimated and the 'error' term e has zero mean and finite variance. Standard errors are clustered at the person-level. [10] The term $\alpha(t)$ is a function of t and completes the baseline logit-hazard model as shown in Eq. (1). For further details on this functional form, see Allison (1984).

Interpretation of the coefficients is as follows: β_1 estimates the impact of the FMLA on the "risk" of giving a birth among ineligible women, β_2 estimates the impact on birth for women who would have been eligible before the FMLA and β_3 is the coefficient of the interaction term and represents the DID estimate. I am primarily interested in the coefficient β_3, which captures the policy impact. It measures the differences in the changes of the hazard estimates of the policy between the treatment and comparison groups. In other words, it shows whether FMLA-eligible women's response to a birth is different from that of ineligible women.

Data

National Longitudinal Survey of Youth 1979 (NLSY79) Cohort

The main dataset for the analysis of fertility outcomes is based on micro data from the 1989 to 2010 waves of the National Longitudinal Survey of Youth 1979 (NLSY79) cohort. Since 1979, interviews have been conducted on 12,686 U.S.-born young men and women who were 14 to 22 years old in 1979 and the respondents have been followed and interviewed annually up until 1994 and biennially from 1996. The NLSY79 is a large nationally-representative survey designed to provide comprehensive information on labor market activities and demographic information. There is an oversampling of Blacks, Hispanics, and economically disadvantaged non-Black/non-Hispanic individuals in the NLSY79. However, the NLSY79 provides customized sampling weights to adjust for this oversampling.

This panel structure of the NLSY79 data offers several advantages over cross-sectional data. For example, the longitudinal nature of the data allows one to follow the same individuals. This is important for this paper in analyzing how women changed fertility behavior over time. A key feature of the NLSY79 in this analysis is the

[10] I also cluster at the state level. Later, I show the results are robust to this specification.

availability of data to determine whether a woman is eligible for FMLA benefits. More specifically, the NLSY79 contains information on hours worked with an employer, whether the respondent has worked for the employer for at least one year as well as specific questions to the respondents on the number of employees employed by their employer.

Over the survey period, it is possible to observe how fertility behavior changed over the years in response to the FMLA policy. By the design of the sample, the youngest person is 24 years old in 1989, the beginning of the analysis period. This means that the youngest person in the sample was 28 years old when the FMLA was implemented. Although the sample may not be representative of fertility in the U.S. during the sample period, almost 65 % of the women over the age of 25 had given birth to at least one child during the same time frame (U.S. Census Bureau 2012, Tables 91).

Dependent Variables and Controls

The dependent variable is a binary indicator that equals one if the woman gave a birth during the survey period 1989 to 2010 and zero otherwise. That is, for each childless woman that is followed each year in the survey period, the variable takes the value of zero until she has a child when it takes the value of one. Specifically, there are two outcome variables, one for whether the person has a first birth and the other is whether the individual has a second birth. In the case of the former, these are women who are childless prior to 1989, while in the case of the latter, I consider only women who never had a second birth before 1989. All the women in the sample have participated in the labor force during the sample period.

For each outcome, I control for a number of characteristics that are likely to affect fertility behavior such as the age, race and ethnicity, education, marital status, cognitive ability in the form of the age-adjusted Armed Forces Qualification Test (AFQT), family income, the number of children a woman had in the past and tastes for children such as women's views on gender roles in the household. I also include work history information such as job tenure. Year dummies are also included to allow for differential characteristics over time. Appendix 1 Table 7 provides the definition of each of the variables used and the units of measurement. I organize the data in a person-year format, resulting in the number of person-years being greater than the actual number of women in the sample.

Measurement and Other Data Issues

There may be drawbacks in interpreting the relationship between the FMLA impact and fertility outcomes. Other influences might bias the results. For example, there is the potential for sorting to occur before the policy, whereby women who desire to have children may choose those jobs that provide maternity leave benefits. Specifically, women with strong fertility desires may be prone to take up jobs that offer maternity leave regardless of the FMLA, while those women with no desire for children may choose firms not covered under the FMLA. In this case, the impact of FMLA on fertility is not clearly identified because it is fertility desires that may be driving women to access leave, as opposed to the availability of leave under the FMLA that is

Table 1 Impact of Fertility Desire on choosing a Maternity Leave Job (OLS Regressions)

Variables	(1)	(2)
	Dependent variable: Employer provides job-protected maternity leave	
Desired number of children	−0.0025 (0.0027)	−0.0014 (0.0035)
Eligible*Desired number of children		−0.0035 (0.0054)
Year dummies	Yes	Yes
State dummies	Yes	Yes
Controls	Yes	Yes
Woman-year observations	46,670	46,670
R-square	0.14	0.16

Eligibility is defined based on the first year of FMLA. Eligible women are those who satisfy all the criteria (see Section 2) to be eligible for FMLA benefits. Ineligible women are those who do not satisfy all the eligibility criteria and who have worked some positive number of hours in each of the years during the sample period 1989–2010. Therefore, women who never worked in any year during the sample period are excluded. Robust standard errors clustered by individual are reported in parentheses. In addition, all regressions control for age, education, marital status, AFQT scores in percentile, race, ethnicity, union status, whether the individual resides in an urban area, family income, number of children ever had, tastes for children as measured by women, gender roles and job tenure

influencing childbearing preferences. Although the study by Averett and Whittington (2001) did not investigate the impact of the FMLA, the authors found no evidence that women who prefer having children were self-selecting into employment providing maternity leave.[11] Selection of this type is unlikely because under the FMLA, there is an implicit waiting period since all FMLA-maternity leave beneficiaries ought to have worked at least one year with the same employer and must have accumulated at least 1,250 h during the past year (see Section 2).

To test the above empirically, Table 1 provides results of some regressions where the dependent variable is a dummy variable that equals one if an individual is employed at a job that offers job-protected maternity leave and zero otherwise.[12] Each specification (columns 1 to 2) includes covariates controlling for various demographic and other characteristics of the person such as age, race and ethnicity and education. The bottom of the table lists all the controls used. I also include state and year fixed-effects.

In column 1, the variable of interest is the "desired number of children" which was asked of each respondent in the 1979 wave of the NLSY79. If women with strong fertility desires are sorting into jobs offering maternity leave then the coefficient on this variable should be positive and significant. The coefficient is negative and not

[11] The authors estimated the effect of the "desired number of children" on the probability of being in a maternity leave job. They found that the coefficient on "desired number of children" was negative and insignificant. A statistically insignificant result indicates little/no evidence of selection bias caused by women sorting into firms offering maternity leave because they want to have children.

[12] Each wave of the NLSY79 asks respondents whether they are employed at a job that offers job-protected maternity leave. Job-protected maternity leave guarantees an individual the right to return to work after being granted leave from the same employer conditional upon the employee spending no more than the maximum allowable leave. As mentioned earlier, varying degrees of job-protected maternity leave existed in states before the FMLA was implemented.

statistically significant; suggesting that desired fertility does not appear to have any impact on the probability of a woman choosing such a job. In another specification (column 2), the variable “desired number of children” is interacted with a dummy variable “Eligible” which equals to one for a woman who would have become eligible for the FMLA in 1993 and zero otherwise. The coefficient on this interaction is positive but is not statistically significant. This is evidence against the possibility that women who might have been eligible for FMLA were already choosing maternity leave jobs in anticipation of the policy. In summary, these results offer evidence to dispel the notion that a widespread number of women with strong fertility desires were sorting into jobs offering maternity leave.[13]

There may also be causality issues where the take-up of maternity leave is itself endogenous. That is, the decision of working women to proceed on maternity leave may be due to child-friendly employment benefits offered in *covered* institutions, but unrelated to FMLA. If eligible women decide to have children because of these benefits, then the impact of the FMLA on fertility outcomes of eligible women is likely to be biased. This potential problem is likely to be minimal because childcare in the U.S. is widely available and is not limited to FMLA-covered institutions. [14]

Other issues surround the use of the NLSY79 data. Because the NLSY79 has only one cohort, one potential issue is that for the women in the sample, the FMLA occurs in the middle of their peak fertility years. For example, for the FMLA to have an effect on the timing of a woman’s first birth, she must not have had a first birth before the age of 28. The characteristics of these women may result in selection bias on the fertility impact of the FMLA. This bias is likely to be downwards because for physiological reasons, older, childless women tend to have a lower probability of having a first birth. It is also possible for an upward bias to occur to the extent that women who are more educated and/or career-oriented may further delay having a first birth, thus making the probability of a first birth more likely at a later age.

There is possible measurement error in deriving the “Eligible” variable. One of the criteria for eligibility for FMLA is that the individual must be employed in a *covered* institution (see Section 2). According to the definition, covered employers comprise all public institutions (including schools) irrespective of the number of employees, private institutions with at least 50 employees, and private elementary and secondary schools regardless of the number of employees. Even though all public schools are covered employers, only some private schools are. Information on the type of private school employer (whether elementary or secondary) is not provided in the NLSY79 and so it is

[13] There is also a selection bias whereby women who desire children may choose to not work. In which case, the probability of giving a birth increases for these ineligible women. By including women who are not working, I am actually underestimating the impact of the policy. Also, women in larger firms with longer tenure and more work hours may be more committed to career and may be less likely to have children or more likely to delay. This would bias against finding fertility effects of FMLA on eligible women.

[14] Indeed, in addition to childcare, employer-side characteristics would be correlated with employment eligibility. As part of the empirical analysis, I consider firm size (as well as income levels) in the investigation of the impact of FMLA eligibility on birth outcomes. In the interest of space, these regression results were not reported.

possible that some private school firms are incorrectly included in the covered category. This discrepancy is likely to be small since the total number of private schools constitutes less than one percent of all firms in the U.S.[15]

In some firms, women receive additional benefits in the form of compensation while on maternity leave. Therefore it may be possible to find two women working in separate firms both of whom are eligible for FMLA but one of them can access paid leave while the other does not. There is no information in the NLSY79 that allows the researcher to differentiate between these two women. If both women decide to have children and proceed on maternity leave, but one did so because of the compensation unrelated to the FMLA, then the impact of FMLA eligibility on the probability of giving birth will be biased. Since only a small fraction of workers enjoy paid family leave benefits this bias is likely to be small.[16]

Empirical Results

Descriptive Statistics

Table 2 shows the summary statistics of the NLSY79 sample of women used in the paper. These are women who had at least some labor market experience over the survey period 1989 to 2010. The first column shows the statistics for all women during the observation period 1989 to 2010. Columns 2 and 3 provide statistics for women who would have been eligible for FMLA benefits before and after the policy was introduced, respectively. Meanwhile, columns 4 and 5 give statistics for ineligible women. These ineligible women are those who have worked some positive number of hours in each of the years of the sample period. More than half of the women are married and this fraction is similar across all categories of women before and after the policy. Across the two groups, the women are similar in age and on average they have acquired about 13 years of schooling. There is a higher proportion of Black eligible women compared to the proportion of Black ineligible women. The opposite is true for women of Hispanic origin.

Non-Parametric Estimation—Kaplan-Meier Estimates

In this section, I use non-parametric estimation analysis to take a first pass at the data. Specifically, I use Kaplan-Meier estimates (Kaplan and Meier 1958) which are considered to be appropriate in samples which are random and where censoring occurs. Kaplan-Meier estimates possess two important characteristics: first, no assumption is

[15] In 2007, there were 33,740 private schools in the US. See the U.S. Department of Education National Center for Education Statistics, Private School Universe Survey 2007–2008 available at http://nces.ed.gov/programs/digest/d09/tables/dt09_058.asp (accessed October 2, 2010). In 2002, the total number of firms as reported by the Census Bureau was 22,974,655. See the U.S. Census Bureau: State and Country Quick Facts available at http://quickfacts.census.gov/qfd/states/00000.html (accessed October 2, 2010).

[16] For example, in 2007, only 8 % of private sector workers could access paid family leave benefits. See U.S. Department of Labor, Bureau of Labor Statistics, *National Compensation Survey: Employee Benefits in Private Industry in the United States* at http://www.bls.gov/ncs/ebs/sp/ebsm0006.pdf (accessed September 21, 2010).

Table 2 Summary Statistics for Women in the NLSY79 sample: Means (standard deviations)

Variables	All women 1989–2010	Eligible women		Working Ineligible women	
		Before (1989–1992)	After (1993–2010)	Before (1989–1992)	After (1993–2010)
	(1)	(2)	(3)	(4)	(5)
Dummy Explanatory Variables					
Eligible	0.378	1.000	1.000	0.000	0.000
Married	0.553	0.510***	0.559	0.566	0.556
Black	0.289	0.307***	0.337	0.239	0.278
Hispanic	0.181	0.154***	0.166	0.181	0.199
Union	0.122	0.167***	0.174	0.053	0.110
Urban	0.760	0.810***	0.751	0.786	0.738
Gender roles:					
A woman's place is in the home	0.145	0.121***	0.115	0.152	0.165
Men should share housework	0.841	0.861***	0.865	0.834	0.825
Women should perform traditional roles	0.259	0.222***	0.216	0.266	0.292
Continuous explanatory variables					
Age	36.387 (6.971)	29.485 (2.484)	39.560 (5.937)	29.451 (2.514)	39.746 (5.904)
Education	13.388 (2.380)	13.614*** (2.231)	14.022 (2.341)	12.901 (2.285)	13.173 (2.403)
Family Income	53.027 (71.610)	46.967*** (89.910)	63.168 (64.213)	42.852 (91.475)	53.580 (56.891)
Parity	1.912 (1.337)	1.623*** (1.242)	1.672 (1.231)	1.956 (1.336)	2.117 (1.385)
AFQT score	43.058 (27.919)	46.455*** (27.943)	46.533 (27.774)	42.111 (27.762)	40.384 (27.743)
Job tenure	271.008 (289.571)	233.627*** (185.904)	424.094 (367.164)	119.768 (141.464)	261.974 (271.409)
Other variables					
High school dropout	0.084	0.060***	0.032	0.132	0.098
High school graduate	0.411	0.388***	0.348	0.445	0.439

Table 2 (continued)

Variables	All women 1989–2010	Eligible women		Working Ineligible women	
		Before (1989–1992)	After (1993–2010)	Before (1989–1992)	After (1993–2010)
	(1)	(2)	(3)	(4)	(5)
Some college	0.277	0.283***	0.310	0.250	0.269
College graduate	0.228	0.269***	0.310	0.173	0.194
Government sector	0.179	0.244***	0.247	0.087	0.163
Private sector	0.639	0.730***	0.562	0.756	0.603
MLS states	0.279	0.273***	0.261	0.299	0.281
Employees	2109.582 (12,596.780)	4457.678** 18637.03)	1467.697 (8464.570)	3701.506 (18243.540)	1053.065 (8324.552)
Hours work	1813.491 (789.393)	2030.388*** (509.000)	2004.759 (695.248)	1483.2279 (815.225)	1786.474 (833.865)
Desired number of children	2.509 (1.492)	2.480* (1.462)	2.478 (1.462)	2.522 (1.505)	2.529 (1.513)
Maternity leave	0.644	0.819***	0.789	0.471	0.586
Covered	0.584	0.890***	0.754	0.345	0.502
N	48,498	6,276	12,040	9,394	20,788

Eligibility is defined based on the first year of FMLA. Eligible women are those who satisfy all the criteria (see Section 2) to be eligible for FMLA benefits. Ineligible women are those who do not satisfy all the eligibility criteria and who have worked some positive number of hours in each of the years during the sample period 1989–2010. Therefore, women who never worked in any year during the sample period are excluded. Statistical levels of significance, based on the difference between means of eligible and ineligible women before FMLA, are as follows: * indicates p <0.10, ** indicates p <0.05, *** indicates p <0.01

made about the specific sample distribution and second, these estimates are consistent for a wide range of distribution classes.

The Kaplan-Meier plots illustrate the survival functions (showing the proportion of women remaining after a birth) for eligible and ineligible women for first and second birth outcomes. These are shown in Appendix Figs. 3 and 4, respectively (similar information is presented in Appendix Tables 10 and 11). There are notable differences for the two categories of females, especially after the FMLA. In Appendix Fig. 3 (see corresponding estimates in columns 2 and 3 of Appendix Tables 10 and 11), the survivor curve for the exit to a first birth for FMLA-eligible women (treatment group) has a much sharper drop after the reform than for other women (control group). After years 1 (1989), 3 (1991), 16 (2004), the probability of "surviving" without having a first birth for an "eligible" woman reduces from 93 % to 79 % then to 54 %, respectively while for ineligible women the respective probabilities are 90 %, 79 % and 61 %. With respect to the likelihood of *surviving* without having a second birth, although the situation is less defined there is still some evidence from Appendix Fig. 4 (see corresponding estimates in columns 5 and 6 of Appendix Tables 10 and 11) suggesting that eligible women are more likely to have a second birth compared to their ineligible counterparts after the FMLA.

While the results are intuitive, these non-parametric estimates may be sensitive in the presence of individual demographic and socioeconomic factors which may vary with the introduction of the FMLA. Inclusion of these characteristics will further help to isolate the effect of the policy. The above analysis has provides a basis to further examine the impact of the policy on fertility behavior using a parametric approach to estimation.

Discrete-Time Hazard Estimates: Probability of having a First and Second Birth

In this section, I use a discrete-time hazard model to present formal estimates of Eq. (1) to test whether the introduction of the FMLA resulted in a differential impact on the likelihood of eligible women giving birth relative to other women. Table 3 presents estimates of the impact on the probability of a first birth. Columns 1 and 2 present the marginal effects based on the logit model (Eq. 1) with time dummies included in all specifications. The key variable of interest is the interaction term FMLA*Eligible which indicates whether the birth response among FMLA-eligible women differs from that of ineligible women. In this sample of women, consideration is given only to those who have never given birth to a first child prior to 1989.

The specifications in columns 1 and 2 differ by only the additional controls included in column 2. In both specifications, the coefficient on the interaction term is positive and statistically significant at conventional levels. These results indicate that the introduction of the FMLA has led to a 2.8 percentage point and 5.2 percentage point increase in the probability that an eligible woman will give a first birth, respectively. Using the results from column 2, this is equivalent to an increase in probability of approximately 1.5 % per annum from the baseline.

Table 4 presents estimates of the impact on the probability of a second birth. [17] The women in this sample consist of those who never experienced a second birth prior to

[17] I also ran the model using completed fertility as the dependent variable. The results were qualitatively similar.

Table 3 Estimates from Logit Regression: Likelihood of a First Birth (Marginal Effects)

Variables	(1)	(2)
	Dependent variable: first birth	
Eligible	−0.0069* (0.0040)	−0.0083*** (0.0021)
FMLA	−0.3115*** (0.1065)	−0.1536** (0.0822)
FMLA*Eligible	0.0281* (0.0151)	0.0519** (0.0247)
Age		0.0176*** (0.0042)
Age squared		−0.0003*** (0.0001)
Education		−0.0003 (0.0005)
Married		0.0355*** (0.0032)
AFQT percentile		−0.0001** (0.0000)
Black		0.0008 (0.0029)
Hispanic		−0.0023 (0.0024)
Union		0.0014 (0.0031)
Urban		−0.0013 (0.0022)
Family income		−0.0000 (0.0000)
Parity		0.0195*** (0.0018)
A woman's place is in the home		−0.0016 (0.0029)
Men should share housework		−0.0035 (0.0029)
Women should perform traditional roles		0.0024 (0.0027)
Job tenure		0.0000 (0.0000)
Job tenure squared		−0.0063 (0.0179)
Year dummies	Yes	Yes
Woman-year observations	11,963	11,963
Log-likelihood	−2589.852	−2032.620
Mean of dependent variable	0.21	0.21

The sample consists of all working women who have never given birth prior to 1989. Robust standard errors clustered by individual are reported in parentheses. Statistical levels of significance are: * indicates p <0.1, ** indicates p <0.05, *** indicates p <0.01

1989 and as such include those women considered in the first birth regressions. The results in this table come from the same model specifications as in Table 3. The coefficient on the interaction term is positive and significant across both specifications as shown in columns 1 and 2. Using the results of column 2, the value of the coefficient indicates that eligible women increase their probability of giving a second birth by about 3.0 percentage points following the introduction of the FMLA (or equivalent to an increase in probability of approximately 0.6 % per annum from the baseline).

The coefficients on a number of independent variables in the model show the expected sign in column 2 of Tables 3 and 4. The likelihood of a first or second birth increases with age. Although the education variable is significant only in Table 4, it is negative; while the measure of cognitive ability (AFQT scores) is negative and significant. Being married increases the probability of having a first or second birth while race, ethnicity, union status and place of residence do not have any significant impact.

Table 4 Estimates from Logit Regression: Likelihood of a Second Birth (Marginal effects)

Variables	(1)	(2)
	Dependent variable: Second birth	
Eligible	−0.0061*** (0.0020)	−0.0030*** (0.0009)
FMLA	−0.1707*** (0.0572)	−0.0521* (0.0296)
FMLA*Eligible	0.0369*** (0.0149)	0.0296* (0.0158)
Age		0.0068*** (0.0015)
Age squared		−0.0001*** (0.0000)
Education		−0.0004* (0.0002)
Married		0.0096*** (0.0012)
AFQT percentile		−0.0000* (0.0000)
Black		0.0015 (0.0012)
Hispanic		0.0007 (0.0011)
Union		−0.0005 (0.0012)
Urban		−0.0009 (0.0009)
Family income		−0.0000 (0.0000)
Parity		0.0093*** (0.0009)
A woman's place is in the home		−0.0004 (0.0013)
Men should share housework		−0.0001 (0.0011)
Women should perform traditional roles		0.0011 (0.0011)
Job tenure		0.0000 (0.0000)
Job tenure squared		−0.0016 (0.0055)
Year dummies	Yes	Yes
Woman-year observations	22,263	22,263
Log-likelihood	−3214.284	−2516.343
Mean of dependent variable	0.27	0.27

The sample consists of all working women who have not had a second child before 1989. Robust standard errors clustered by individual are reported in parentheses. Statistical levels of significance are: * indicates p <0.1, ** indicates p <0.05, *** indicates p <0.01

Heterogeneity Across Sector of Employment, Race and Ethnicity and Education

Given the universal nature of the FMLA, it is particularly interesting to investigate what effect the FMLA and its eligibility criteria had on first birth among specific sub-groups of the sample. First, I estimate the effect separately for the government sector and all other sectors using Eq. 1. The coefficient on the interaction term is positive and statistically significant in other sectors but not for those in the government sector. This result is consistent with the fact that all public institutions are *covered* employers, therefore making it far easier for an otherwise ineligible woman to become eligible.[18]

[18] See Section 2 for a definition of a covered employer and how this relates to a woman being eligible for FMLA benefits. I also present models where the sample is disaggregated according to firm size. For firms with less than 50 employees, the coefficient of interest was not significantly different from zero. This result is not surprising because firms in this size category (with a few exceptions as noted in Section 2) by definition are not *covered* institutions, thus making the employees ineligible for FMLA benefits.

Next, I divide the sample by race and ethnicity, specifically Blacks, Hispanics and Whites. The only significant result is obtained for Whites, suggesting that among whites, eligible women increase their probability of giving a first or second birth. One plausible explanation for this outcome: to the extent that whites are more likely to be married and have high income-earning husbands, it makes it more feasible for them to take unpaid leave to have children under the FMLA. In terms of education level, eligible women among those with at least some college experience increase the probability of giving a first or second birth. This particular outcome might be associated with the fact that white eligible women who are utilizing the leave benefits under FMLA the most also happen to be the more educated.[19] This result is also consistent with the findings that college-educated females were more likely to take unpaid maternity leave, while the FMLA policy appeared to have no significant impact on women who had never gone to college (Han, Ruhm and Waldfogel 2007).

Although the results are positive across Blacks and Hispanics, as well as across lower educational levels, the policy appears to lose statistical power and estimates are no longer statistically significant. In the interest of space, the results analyzed in this section are not reported but available upon request. [20]

Sensitivity Analysis

Some states had maternity leave statutes (MLS) prior to the introduction of the FMLA. It is possible individuals residing in states that already had some form of MLS are unlikely to be affected as much by the FMLA, since in some cases, the benefits provided under the FMLA may have been less generous. Also, in 2002 California introduced the Paid Family Leave Bill that provides disability compensation for employees unable to perform work duties due to the birth of a child along with other provisions. In light of these circumstances, I estimate the main model (Eq. 1) by excluding California. In a separate regression, I exclude all states that had an MLS (California is included as well) prior to the FMLA. The results are shown in Table 5 (columns 2 and 3, respectively). Panels 1 and 2 present the results when the dependent variable is the probability of a first birth and second birth, respectively. Even with these exclusions, there is still evidence of an impact of the program on births to eligible women.

Next, a dummy variable equaling one was created for all states that had MLSs and zero otherwise. This dummy variable was interacted with eligibility status. This would give some variation in access to leave that did not turn on in 1993. The results as shown in column 4 (both panels), are robust to this specification. In column 5, I re-estimate the model with a placebo effect. Here, I assume that the policy started in some period other than the year it actually took place. The estimates of the placebo effects can be

[19] Therefore, the more educated are better able to utilize the leave benefits of the FMLA because of better knowledge about policy. Recall, only 58.2% of employees at covered worksites had any knowledge of FMLA (Department of Labor 2007). Another reason why the FMLA has been ineffective in for other groups such as those in marginal employment can be explained by the fact that the majority of these individuals are not covered under the current FMLA eligibility criteria. That is, they are employed in institutions with less than 50 employees. See further details in the section on Sensitivity Analysis.

[20] In addition, I run regressions based on the size of the firm and find that the effect of the interaction term (FMLA*eligibility) is significant at the 10% level for large firms (those with at least 500 employees).

Table 5 Estimates from Logit Regression: Sensitivity Analysis on the likelihood of a Birth (Marginal Effects)

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
PANEL 1: Dependent variable—First birth	Full sample	Excluding California	Excluding MLS states	Including Eligible * MLS interaction term	Placebo Effect: FMLA lasted 1990–1992†	Linear time trend interacted with eligibility (OLS)	Eligibility based on firm size (hours worked can vary)	Eligibility based on hours worked (firm size can vary)
FMLA*Eligible	0.0519** (0.0247)	0.0502** (0.0246)	0.0426* (0.0240)	0.0458* (0.0240)	−0.0231 (0.0222)	0.0375*** (0.0121)	0.0568** (0.0258)	0.0642**(0.0307)
Woman-year observations	11,963	10,576	8,157	11,963	11,963	11,963	11,963	11,963
Log-likelihood/ R-square	−2032.620	−1899.793	−1497.304	−2035.823	−2045.534	0.12	−2036.160	−2039.747
PANEL 2: Dependent variable—Second birth	Full sample	Excluding California	Excluding MLS states	Including Eligible * MLS interaction term	Placebo Effect: FMLA lasted 1990–1992†	Linear time trend interacted with eligibility (OLS)	Eligibility based on firm size (hours worked can vary)	Eligibility based on hours worked (firm size can vary)
FMLA*Eligible	0.0296* (0.0158)	0.0318* (0.0183)	0.020 (0.0163)	0.0162 (0.0102)	−0.0009 (0.0071)	0.0180** (0.0077)	0.0315** (0.0164)	0.0463** (0.0209)
Woman-year observations	22,263	19,842	15,796	22,263	22,263	22,263	22,263	22,263
Log-likelihood/ R-square	−2516.343	−2310.523	−1796.212	−2514.523	−2511.297	0.08	−2518.706	−2508.081
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors clustered by individual are reported in parentheses. Statistical levels of significance are: * indicates p <0.1, ** indicates p <0.05, *** indicates p <0.01. † indicates p=0.00 for the Likelihood-ratio test on the difference between the placebo and the true results (column 1). In addition, the AIC/BIC provide strong support for the true model. Controls used in the regressions are age, education, marital status, AFQT scores in percentile, race, ethnicity, union status, whether the individual resides in an urban area, family income, number of children ever had, tastes for children as measured by women, gender roles and job tenure compared with the estimates of the policy effect in column 1. As expected, the placebo effects are not statistically significant, thus providing further support to the identification strategy. [21] In column 6, I consider the possibility that there may be some pre-existing differential trends in fertility between eligible and non-eligible women. To control for this, I include a linear time trend and interacted it with the "Eligibility" dummy. Once again, the results are robust to this specification. [22]

In the next set of robustness checks, I assume that the criteria for eligibility is based either on firm size, with hours worked allowed to vary (column 7) or eligibility is based solely on hours worked with the size of the firm allowed to vary (column 8). Recall in Section 2, for the purposes of the FMLA, eligibility is based on a minimum number of hours worked as well as being employed in a firm of with a minimum size (i.e. at least 50 employees). Compared with the base estimates in column 1, the results in columns 7 and 8 are larger in the majority of the cases. Even more interestingly are those estimates in column 8, where the coefficient is almost one and a half times as large as in the case of a first birth. This is interesting because the results can have important policy implications. One interpretation of this finding is as follows: if the eligibility criteria for receiving FMLA benefits were based only on the amount of hours an employee works (disregarding the size of the firm), then the fertility impact could be larger. Given that almost 90 % of the firms in the U.S. employ 20 or less individuals, it would appear that a change in policy that relaxes the eligibility criteria can have a greater fertility impact through its effect on a broader spectrum of the labor force.[23]

Other robustness checks reported are as follows: clustering standard errors at the state-level, regressions based on marital status, models including a dummy variable to account for states that had maternity leave statutes prior to the FMLA and models where eligibility for FMLA is based on satisfying the criteria in each survey year. These results are shown in Appendix Table 9.

While it is not possible to guarantee that the FMLA was the sole cause of the relative increase in fertility behavior among eligible women, there is no evidence that some other factor(s) with a similarly universal impact and unrelated to the FMLA could have affected eligible women to the same degree as the previous findings indicate.

Parametric Estimation using Log-normal Regressions

The results obtained above have been generated from a set of semi-parametric models. These models are more appropriate in cases where the distribution of the outcome is not exactly known. However, there are some parametric (distribution-based) models prevalent in the social sciences and that are appropriate in this analysis. One commonly–used model is the log-normal model associated with skewed outcomes and is also used to measure the response times of an individual to an event.[24] The model measures the log of survival time. Specifically, the log-normal model assumes a log-normal distribution of the form:

$$\log(\mathrm{t}_i) = \beta_1 \mathrm{FMLA}_t + \beta_2 \mathrm{Eligible}_i + \beta_3 (\mathrm{FMLA}_t * \mathrm{Eligible}_i) + \beta_4 \mathbf{X}_{it} + \varepsilon_{it} \quad (2)$$

[21] The p-value for the Likelihood Ratio test on the difference between the placebo and the true results is zero in both panels. In addition, the AIC/BIC provide strong support for the true model.

[22] I also run the same regression, this time including a quadratic trend and interacting with eligibility status. The results (not shown) are also robust to this specification.

[23] Data source: http://www.census.gov/epcd/www/smallbus.html (accessed January 25, 2011).

[24] Inspection of the data reveals positive skewness in the outcomes used in this analysis.

where the dependent variable measures the log of survival time in years, t. The remainder of the model is the same as in Eq. 1. The error term is assumed normal with zero mean and constant variance.

Table 6 presents the results of the log-normal time-to-birth regressions for the two birth outcomes. The specifications in columns 1 and 2 are similar to the models shown in columns 2 of Tables 3 and 4, respectively. The sign of the coefficient on the FMLA*Eligible interaction term in columns 1 and 2 of Table 6 is opposite those of the discrete-time hazard models in columns 2 of Tables 3 and 4, respectively. Since eligible women since the FMLA have a *greater* hazard of giving birth then it implies that these women will take *less* time on average to do so. In both outcomes, the coefficients are statistically significant. For an interpretation of the coefficient on the interaction term in Table 6: following the FMLA, eligible women are taking on average 44.1 % less time before having a first birth (column 1) while taking 30.1 % less time before having a second birth (column 2).

Table 6 Estimates from Log-normal Regression: Time-to-birth in years

Variables	(1) Dependent variable: log of Time to birth in years	(2)
	First birth	Second birth
Eligible	0.158** (0.063)	0.147** (0.058)
FMLA	0.906*** (0.086)	0.822*** (0.075)
FMLA*Eligible	−0.441*** (0.107)	−0.301*** (0.092)
Age	−0.621*** (0.085)	−0.510*** (0.089)
Age squared	0.011*** (0.001)	0.009*** (0.001)
Education	0.017 (0.013)	0.025** (0.011)
Married	−0.683*** (0.058)	−0.529*** (0.051)
AFQT percentile	0.002 (0.001)	0.003** (0.001)
Black	−0.020 (0.077)	−0.034 (0.067)
Hispanic	0.021 (0.081)	−0.002 (0.072)
Union	−0.015 (0.077)	0.001 (0.075)
Urban	0.019 (0.067)	−0.001 (0.059)
Family income	−0.000 (0.000)	0.000 (0.000)
Parity	−0.551*** (0.029)	−0.644*** (0.032)
A woman's place is in the home (Gender role1)	0.074 (0.078)	0.017 (0.074)
Men should share housework (Gender role2)	0.053 (0.073)	−0.064 (0.067)
Women should perform traditional roles (Gender role3)	−0.028 (0.065)	−0.002 (0.059)
Job tenure	0.000 (0.000)	−0.001** (0.000)
Job tenure squared	−0.124 (0.465)	0.356 (0.361)
Woman-year observations	11,963	22,263
Log-likelihood	−1353.655	−1732.872

Robust standard errors clustered by individual are reported in parentheses. Statistical levels of significance are: * indicates p <0.1, ** indicates p <0.05, *** indicates p <0.01

To put this into context, Appendix 1 Table 8 presents some calculations on the time taken for eligible women to give a first and second birth before and after the FMLA. Column 1 gives the baseline average waiting time in years. In the sample, eligible women took just over two years before giving a first or second birth. With the introduction of the FMLA, they have reduced this waiting time by over a year for a first birth and by 0.71 years (approximately 8.5 months) for a second birth, resulting in a fall in the average waiting time before giving a birth of 0.86 years (or an equivalent of 10 months) as shown in column 2. The predicted new waiting times are approximately 1.3 years for giving a first birth and just under two years for a second birth (column 3). On average, the waiting time has been reduced to 1.5 years.

To illustrate how the probability of giving a birth changed over the years, Figs. 1 and 2 plot the fitted log-normal survival functions between FMLA-eligible women and ineligible women for the first and second birth outcomes. The plotted predicted survival functions clearly indicate eligible women have lower probabilities of not having either a first or a second birth. The differences in the probabilities between the two groups of women get larger after about the fourth year which is when the FMLA began.

Discussion and Conclusion

Timing Versus Completed Fertility

The hazard models tell us about the timing of events. Indeed, the results suggest that eligible women post-FMLA are giving birth to their first child a year earlier and their second child 8.5 months earlier. From the survival graphs, it can be inferred that completed fertility is higher for eligible women compared to those not eligible for FMLA. In fact, when I ran models using completed fertility as the dependent variable,

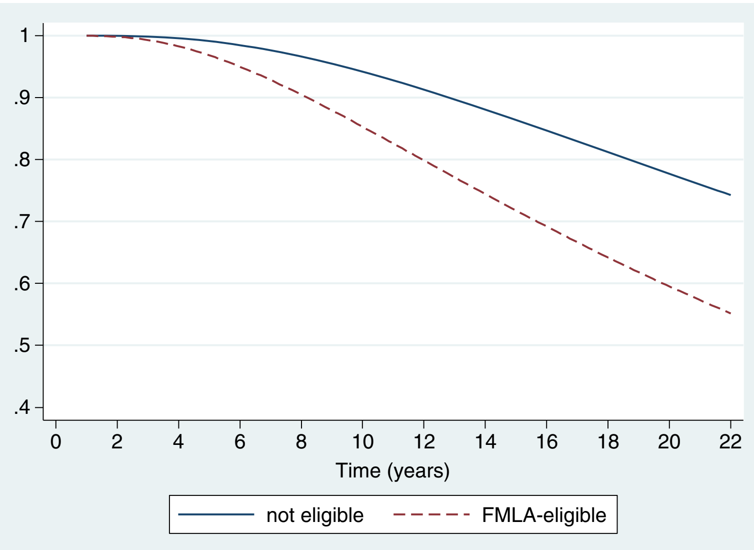

Fig. 1 Log-normal Survival Plots of not giving a first birth

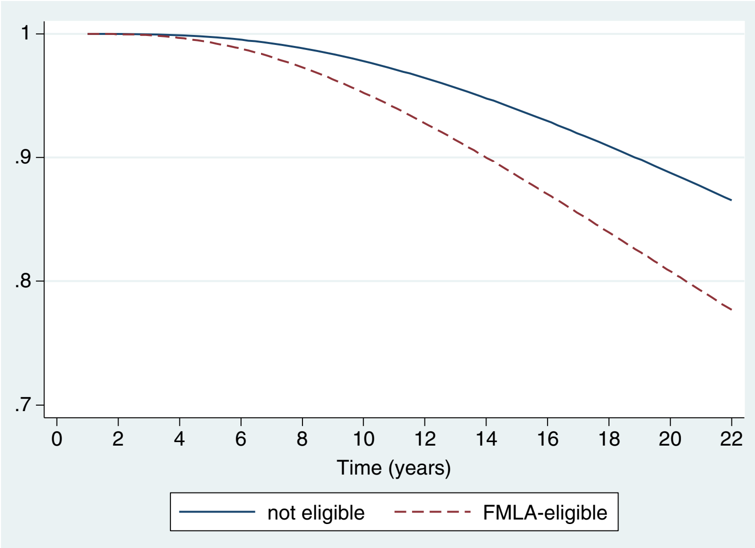

Fig. 2 Log-normal Survival Plots of not giving a second birth

the results were qualitatively similar. In these results, since the youngest women are 45 years at the end of the sample period, there is likely some bias from those women yet to complete their fertility and who might end up having more children.

Important for this analysis is the overall impact on fertility behavior—the main question of this research. In this regard, three relevant questions arise: Is the effect strictly a timing effect? Is there also an effect on eventual childlessness or completed fertility as well? Do the policy implications of the FMLA differ if the effect is only timing versus whether it affects completed fertility? Recent US fertility data show that there has been an increasing *postponement effect*—a decrease in completed fertility as result of timing considerations.[25] Specifically, women are further delaying the onset of childbearing. This is particularly true for the younger cohort of women (under 25) who are likely pursuing economic activities (such as employment) and education in preference to children. Simultaneously, there has been an increase in fertility rates among older women. Since the US fertility rates have hovered around 2 births per woman over the last decade, one conclusion is that completed fertility has remained largely unaffected. The NLSY sample in this study goes through 2010 where the youngest respondents are 45 in 2010, so these individuals have almost completed their fertility cycle at an average age of 51. As a result, the effects observed might be more suggestive of timing and possibly understating the implications of the FMLA in terms of its effect on completed fertility for younger cohorts. Further, the FMLA only affects a relatively small portion of the working population and on the basis of the previous results, given a more expanded form of the FMLA, there are policy implications relating to potentially speeding up the timing of births and thereby affecting completed fertility.

[25] See the 2011 Statistical Abstract of U.S. Census Bureau at http://www.census.gov/compendia/statab/cats/births_deaths_marriages_divorces.html (Accessed July 19, 2011).

Conclusion

This paper examines the impact on fertility outcomes of a policy—the Family and Medical Leave Act (FMLA)—that allows women to proceed on job-protected maternity leave. Because the FMLA has clearly-defined eligibility criteria and empirical evidence suggests that it may have been associated with an increase in work leave taken, it is possible to identify the impact on fertility by comparing the outcome among eligible women (who are able to access FMLA leave) with the fertility outcomes of those who are not eligible for FMLA.

I find that the implementation of the FMLA has resulted in eligible women increasing their probabilities of giving birth to a first and second child. The magnitude of the effects appears larger for a first birth. Specifically, among eligible women, the FMLA increased the probability of giving birth by about 5 percentage points for a first birth and 3 percentage points for a second birth. These changes are equivalent to respective increases in probabilities of approximately 1.8 and 1.1 % per annum. The results of the analysis also show that eligible women are giving birth to the first child a year earlier and about 8 months earlier for the second child. I also consider the impact of the policy across sectors, race and ethnicity, and education level. These results indicate that the FMLA is more effective in non-government sectors. Among Whites, eligible females have significantly higher probabilities of giving a first or second birth. Meanwhile, eligible women with at least some education experience at the college level are more likely to give a first or second birth since the FMLA.

These results are consistent with increased leave being taken by eligible recipients (Ross 1998; Waldfogel 1999; Han and Waldfogel 2003; Han, Ruhm and Waldfogel 2007). The results are also consistent with the findings that college-educated females were more likely to take unpaid maternity leave while the policy appeared to have no significant impact on women who had never gone to college (Han, Ruhm and Waldfogel 2007). In addition to the effectiveness of the policy in influencing child preferences, by allowing eligible women the right to return to their former jobs after giving birth, the FMLA has effectively improved the labor outcomes of new parents (Waldfogel 1999). Beyond these, the findings raise the possibility that the policy may have improved other outcomes not yet explored. For instance, the long-term educational attainment and health outcomes of children born to FMLA-eligible women represent important areas for future research.

These findings may have important policy implications, especially when compared to the efficacy of more generous policies in Western European countries such as Italy, Spain and Germany. In these countries, the emphasis is on addressing falling and low fertility. Yet, despite the presence of significantly long leave periods, some of which is paid, the fertility impact of these polices remains ambiguous.

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Appendix

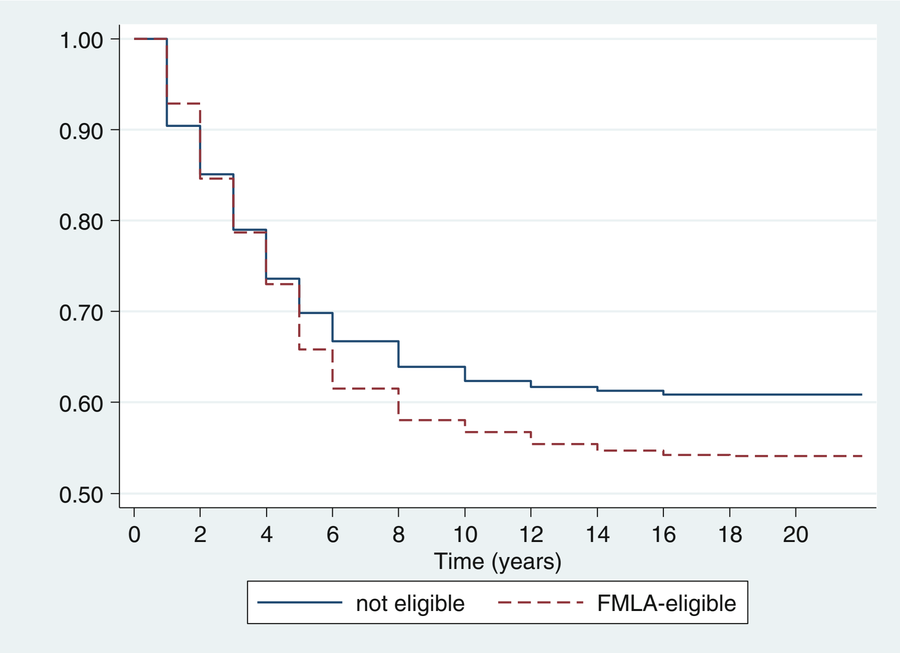

Fig. 3 Kaplan-Meier Survival Function of the proportion of women remaining after a first birth by year and eligibility status

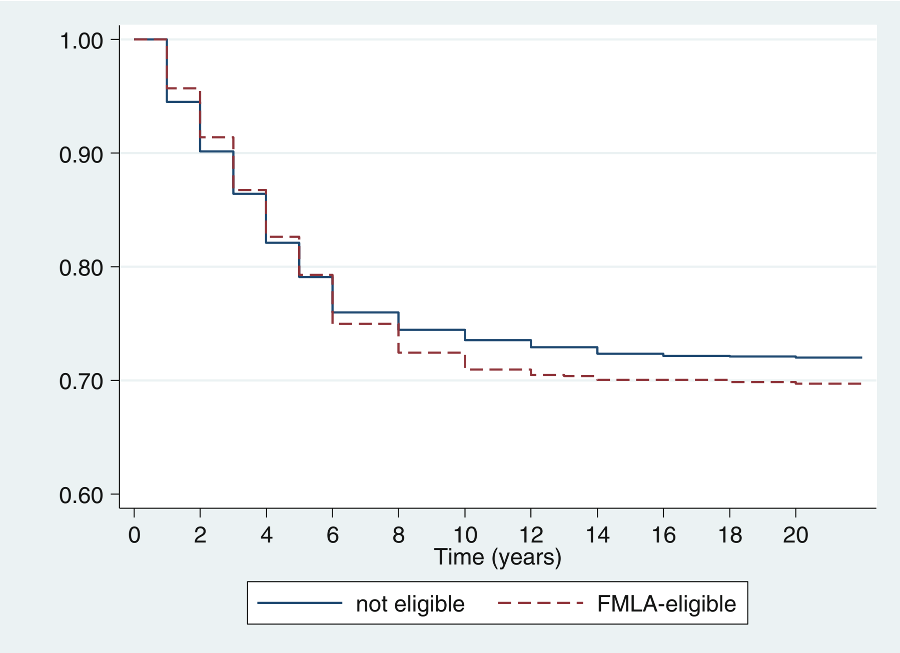

Fig. 4 Kaplan-Meier Survival Function of the proportion of women remaining after a second birth by year and eligibility status

Table 7 Description of Variables

Variables	Variable Definition
Outcome variables	
First birth	Equals one if first child is born on or after 1989 and zero otherwise
Second birth	Equals one if second child is born on or after 1989 and zero otherwise
Dummy explanatory variables	
FMLA	Equals in a year the FMLA is in effect, zero otherwise
Eligible	Equals one if individual satisfies the eligibility criteria for obtaining benefits under the FMLA, zero otherwise
Married	Equals one if married, zero otherwise
Black	Equals one if Black, zero otherwise
Hispanic	Equals one if Hispanic, zero otherwise
Union	Equals one if part of a union in primary job, zero otherwise
Urban	Equals one if current residence is urban, zero otherwise
A woman's place is in the home	Equals one if respondent thinks a woman's place is in the home, zero otherwise
Men should share housework	Equals one if respondent thinks men should share housework, zero otherwise
Women should perform traditional roles	Equals one if respondent thinks women should perform traditional roles, zero otherwise
Continuous explanatory variables	
Age	Age in years
Education	Years of schooling (Highest grade completed)
Family income	Total net family income in thousands of dollars
Parity	Number of children ever born
AFQT score	Armed Forces Qualifying Test (AFQT) score in percentile
Job tenure	Total tenure (in weeks) with employer at primary job
Other variables	
High school dropout	Equals one if highest grade completed is less than 12, zero otherwise
High school graduate	Equals one if highest grade completed is 12, zero otherwise
Some college	Equals one if highest grade completed ranges from 13 and 15, zero otherwise
College graduate	Equals one if highest grade completed is at least 16, zero otherwise
Employed	Equals one if respondent is currently employed, zero otherwise
Government sector	Equals one if employed with the state or federal government, zero otherwise
Private sector	Equals one if employed with a private-owned organization, zero otherwise
MLS states	Equals one if state is had a job-protected Maternity Leave statute (MLS) prior to the FMLA, zero otherwise
Employees	Number of employees at respondent's current job location
Hours work	Number of hours worked in the past calendar year
Number of children desired	Number of children desired
Maternity leave	Equals one if employer offers job-protected maternity leave, zero otherwise
Covered	Equals one if individuals works in a covered institution, zero otherwise

Table 8 Predicted response time (years) before getting a birth (Eligible Women)

	(1) Eligible Women	(2)	(3)=(1)+(2)
	Before FMLA (Baseline)	Marginal effect Units (%)	After FMLA
First birth	2.32	−1.02 (−44.1 %)	1.30
Second birth	2.37	−0.71 (−30.1 %)	1.66
Simple Average	2.35	−0.86	1.48

Probabilities were calculated using the coefficients presented in columns 1 and 2 of Table 6 for a first birth and second birth, respectively. All evaluations are done at the sample means

Table 9 Estimates from Logit Regression: Sensitivity Analysis on the likelihood of a Birth (Marginal Effects)

Variables	(1)	(2)	(3)	(4)	(5)	(6)
PANEL 1: Dependent variable—First birth	Full sample	Full sample‡	Married	Never Married and Other	Include MLS dummy	Eligibility defined in each year
FMLA*Eligible	0.0519** (0.0247)	0.0519** (0.0224)	0.0741*** (0.0276)	0.0158 (0.0143)	0.0513** (0.0234)	0.0596** (0.0285)
Woman-year observations	11,963	11,850	4,767	7,196	11,963	11,963
Log-likelihood/ R-square	−2032.620	−2134.424	−1429.874	−562.991	−2155.511	−2037.542
PANEL 2: Dependent variable—Second birth	Full sample	Full sample‡	Married	Never Married and Other	Include MLS dummy	Eligibility defined in each year
FMLA*Eligible	0.0296* (0.0158)	0.0296** (0.0166)	0.0253 (0.0178)	0.0311 (0.0366)	0.0332* (0.0181)	0.0422* (0.0218)
Woman-year observations	22,263	22,096	10,782	11,481	22,263	22,263
Log-likelihood/ R-square	−2516.343	−2574.909	−1882.723	−1882.723	−2587.212	−2,512-491
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes	Yes

Robust standard errors clustered at the individual level are reported in parentheses, with the exception of column 2, where ‡ indicates standard errors are clustered at the state level. Controls used in the regressions are age, education, marital status, AFQT scores in percentile, race, ethnicity, union status, whether the individual resides in an urban area, family income, number of children ever had, tastes for children as measured by women, gender roles and job tenure. Statistical levels of significance are: * indicates $p<0.1$, ** indicates $p<0.05$, *** indicates $p<0.01$

Table 10 Kaplan-Meier Survival Function Estimates of the proportion of women remaining after a birth outcome

Year	First birth			Second birth		
	All women (1)	FMLA-eligible women (2)	Non-eligible women (3)	All women (4)	FMLA-eligible women (5)	Non-eligible women (6)
1 (1989)	0.916	0.929	0.904	0.950	0.957	0.945
2 (1990)	0.849	0.846	0.851	0.907	0.914	0.902
3 (1991)	0.789	0.787	0.790	0.866	0.868	0.864
4 (1992)	0.734	0.730	0.736	0.823	0.827	0.821
5 (1993)	0.680	0.659	0.699	0.792	0.793	0.791
6 (1994)	0.643	0.615	0.667	0.756	0.750	0.760
8 (1996)	0.612	0.581	0.639	0.736	0.724	0.744
10 (1998)	0.598	0.567	0.624	0.725	0.710	0.735
12 (2000)	0.588	0.555	0.617	0.719	0.705	0.729
13 (2001)	0.588	0.555	0.617	0.719	0.704	0.729
14 (2002)	0.583	0.547	0.613	0.714	0.701	0.723
16 (2004)	0.578	0.543	0.609	0.713	0.701	0.722
17 (2005)	0.578	0.543	0.609	0.713	0.701	0.722
18 (2006)	0.577	0.541	0.609	0.712	0.698	0.721
19 (2007)	0.577	0.541	0.609	0.712	0.698	0.721
20 (2008)	0.577	0.541	0.609	0.711	0.697	0.720
21 (2009)	0.577	0.541	0.609	0.711	0.697	0.720
22 (2010)	0.577	0.541	0.609	0.711	0.697	0.720

Table 11 Kaplan-Meier Survival Function Estimates of the number of women remaining after a birth outcome

Year	First birth			Second birth		
	All women (1)	FMLA-eligible women (2)	Non-eligible women (3)	All women (4)	FMLA-eligible women (5)	Non-eligible women (6)
1 (1989)	1991	915	1076	3237	1,306	1,931
2 (1990)	1,797	842	955	3,019	1,236	1,783
3 (1991)	1,450	658	792	2,501	1,009	1,492
4 (1992)	1,331	608	723	2,360	952	1,408
5 (1993)	1,226	561	665	2,228	904	1,324
6 (1994)	1,124	504	620	2,112	862	1,250
8 (1996)	1,036	462	574	1,965	797	1,168
10 (1998)	952	427	525	1,856	753	1,103
12 (2000)	892	403	489	1,756	714	1,042
13 (2001)	837	376	489	1,673	683	1,042

Table 11 (continued)

Year	First birth			Second birth		
	All women (1)	FMLA-eligible women (2)	Non-eligible women (3)	All women (4)	FMLA-eligible women (5)	Non-eligible women (6)
14 (2002)	836	375	461	1,670	680	990
16 (2004)	790	355	435	1,593	653	940
17 (2005)	790	355	435	1,518	653	885
18 (2006)	746	338	408	1,517	633	884
19 (2007)	717	327	408	1,451	613	884
20 (2008)	716	326	390	1,450	612	838
21 (2009)	643	296	347	1,289	557	732
22 (2010)	623	285	338	1,127	492	635

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