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Female Labour Supply and Household Employment Shocks: Maternity Leave as an Insurance Mechanism

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Abstract

This paper investigates the role of unpaid maternity leave in providing household insurance against paternal employment shocks. The main outcome is the timing of a mothers' return to work after having a child. Exploiting the US Family and Medical Leave Act, we find that mothers eligible for maternity leave speed up their return to work in response to a paternal shock, with the conditional probability of being in work 49% higher than in households with no unpaid maternity leave. Further evidence is provided on the insurance role of unpaid maternity leave through i) no significant interaction between *paid* maternity leave and the paternal shock and ii) smoothing of consumption effects of the shock for households covered by unpaid leave.

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Key Words: Female Labour Supply; Insurance; Maternity Leave. JEL Codes I30; J13; J20; J64.

1 Introduction

An important policy question asks how families insure themselves against shocks to income or employment. We know there is imperfect insurance as both consumption and child human capital respond to unexpected changes to income.[1] Since women have entered the labour force, female labour supply has become a potential form of household insurance.[2] However despite this, insurance is imperfect and there are welfare implications to household shocks.

This paper analyses whether access to unpaid maternity leave offers an insurance role by increasing mothers' responsiveness to paternal employment shocks. Whilst the benefits of maternity leave on female labour supply[3] and child outcomes[4] have been examined, this paper draws upon a third benefit which is as yet unstudied - the insurance role of maternity leave. A mother who is eligible for unpaid maternity leave has a right to return to work after the birth. If her partner loses a job around the timing of the birth, the right to work reduces search frictions and makes it easier for her to smooth the effect of the job loss.

We exploit time-state variation across US states in the implementation of unpaid maternity employment protection. In the US there was no federal legislation regarding maternity leave until the Family and Medical Leave Act (FMLA) was introduced in 1993, which allowed 12 weeks of unpaid maternity leave.[5] However, some states implemented their own version of the policy as early as 1972[6]. Although the FMLA is 20 years old the implications of this paper reach beyond an analysis of the policy itself, by informing about the mechanisms

[1] Attanastio and Davis (1996) reject full insurance of consumption against shocks and Blundell et al. (2008) find only partial insurance of consumption against permanent shocks and full insurance of transitory shocks for non-poor households. Carneiro et al. (2010) estimate human capital responses to permanent income shocks which decline across the child life cycle and responses to transitory shocks which are flat across child age. Finally, Carneiro and Ginja (2015) argue that parental investments in child human capital are close to being fully insured, with only small response of investments to permanent shocks and full insurance against transitory shocks.

[2] See for example Blundell et al. (2016).

[3] for example Waldfogel (1999), Berger and Waldfogel (2004), Hofferth and Curtin (2006), Lalive and Zweimuller (2009), Lalive et al. (2014), Schönberg and Ludsteck (2014).

[4] see Rhum (2000), Gregg and Waldfogel (2005), Gregg et al. (2005) Baker and Mulligan (2008a, 2008b), Liu and Skans (2010), Rasmussen (2010), Carneiro et al. (2015a), Rossin (2011), and Dustmann and Schönberg (2012).

[5] Conditions for eligibility, discussed in Section 2, include working for the employer for at least 1250 hours in the year before birth and a firm size of at least 50.

[6] Waldfogel (1999) found leave to increase as a result of FMLA, Berger and Waldfogel (2004) found for mothers working before birth, those covered by FMLA were more likely to take at least 12 weeks and Hofferth and Curtin (2006) found FMLA to raise employment post childbirth but lower wages.

households use to insure against shocks.

A difference-in-difference approach identifies the insurance mechanism through an interaction of layoff and FMLA. The monthly labour market status of mothers and fathers is constructed using the US Panel Study of Income Dynamics (PSID), starting from the month they have a child. The parents are followed up to the time that the mother shifts labour state[7], either to re-enter the labour market or to have another child. In between the birth and the change of labour state the father may experience an exogenous shock by losing his job.[8] This is a meaningful shock as 8% of the sample experience layoff prior to the mother shifting state. Using a duration model, we estimate whether mothers speed up their return to work after birth in response to the paternal job loss and specifically whether the marginal effect of layoff is heterogeneous by eligibility to unpaid maternity leave. In considering a mother's decision to return to work, we control for her future fertility decisions[9] using a competing risk methodology.

We find the conditional probability of observing a mother in work after a paternal shock is 49% higher in households with employment protection around childbirth, relative to a household with no paternal shock. This suggests that mothers with no maternity employment protection are less able to use their labour supply to insure households. The results are statistically significant for movements into full- but not part-time work which is intuitive as FMLA only offered employment protection if mothers had worked 1250 hours in the previous year.

To give further evidence that this responsiveness of the return to work is due to the insurance role of unpaid leave, we repeat analysis focusing on whether mothers exposed to a paternal shock speed up the return to work if eligible for paid maternity leave. The intuition is that there is less financial benefit of returning to work early whilst on paid leave and indeed, we find no significant interaction effect of a layoff with eligibility to paid leave. Finally, using data on annual food consumption, we find that whilst a layoff lowers household

[7]the destination state of an individual (to remain at home, have a child or move to work) will be referred to as the labour state.

[8]Similarly to Rhum (1991) and Stevens (1997) the job loss is recorded as exogenous if the reason was recorded as plant closure, laid off or being fired.

[9]For example, Del Bono et al. (2012) find fertility effects of female job displacement.

food consumption, this effect is mitigated if households are covered by FMLA, evidence of smoothing of food consumption through FMLA.

Employment protection through FMLA provides insurance for the mother from losing *her* job whilst taking some time off after birth. The additional insurance role studied in this paper is insurance against *paternal* shocks by elimination of search frictions. A large literature models labour market participation in the presence of search frictions (see Mortensen and Pissarides 1999 and Mortensen 2011 for a review). Whilst classically the model consists of two labour market states of employment or unemployment, a number of papers have added the state of non-participation, which is distinguished from unemployment through passive job search behaviour, see Kim (2001), Garibaldi and Wasmer (2005), Yip (2003), Pries and Rogerson (2009) and Moon (2011). In particular, Pries and Rogerson (2009) describe labour market frictions as a fixed cost which make job search more costly and note that "Increases in this fixed cost make non-participation more attractive at the margin." (Pries and Rogerson 2009, p. 569). An increase in this fixed cost is analogous to a limit in employment protection after birth. Section 5.4 explores heterogenetiy in the insurance effect of unpaid leave, by three variables which typically proxy for search frictions - the business cycle, local labour market conditions and maternal education..

The paper is related to a literature which has found that female labour supply as an insurance mechanism is responsive to the level of her partner's unemployment insurance (Cullen and Gruber 2000), health insurance (Buchmueller and Valletta, 1999) and Medicaid (Winkler, 1991). Our paper instead links the female labour supply response to a paternal shock across eligibility to unpaid maternity leave.

The identification comes from the exogenous paternal employment shock and we show our results are robust to two potential sources of endogeneity of the shock - predictability of the event through past experience of layoff and through anticipation effects.

There are important policy implications from this paper. If mothers are less able to insure their households against paternal shocks, the welfare consequences will be felt by adults (Black et al. 2015 find health effects in Norway) and children (see Duncan and Brooks-Gunn 1997, Carneiro and Heckman 2003, Currie 2009, Carneiro et al. 2010, Dahl and Lochner 2012 and Carneiro et al. 2015b for examples). However the other side of the coin is the evidence which suggests negative consequences for child development of early maternal return to work (within the first 12 weeks), through lowering immunizations and breast-feeding, worsening child behavioural problems (Berger et al. 2005) and cognitive outcomes (Baum 2003a). Moreover, combined with the evidence in the paper of no significant movement into work in response to the layoff if mothers are covered by paid leave, this suggests in an extension of the FMLA to offer paid maternity leave in the first crucial months of childhood could allow mothers to insure household income without compromising child development.

The paper proceeds as follows. Section 2 describes the maternity leave legislation in the US. Section 3 details the data and section 4 the methodology. Results are in section 5 and section 6 assesses the insurance role of FMLA. Finally section 6 concludes.

2 Maternity Leave Legislation

Whilst the FMLA policy was implemented over 20 years ago, a debate continues to the present day about the need for an extension to put US maternity leave policy more in line with other OECD countries. The entitlement to weeks of leave around childbirth in the US is currently the lowest in all OECD countries.[10] US federal legislation grants mothers 12 weeks of unpaid leave around the birth of a child but the US is the only OECD country to offer no paid maternity leave. Up until 1993 there was no federal legislation regarding maternity leave in the US at all. Despite this, some states had chosen to implement a job protected maternity leave.[11] Mothers without access to maternity employment protection could accrue annual leave, to spend some time at home with their new child. In 1993 legislation was passed and now employers in all states are obliged to offer 12 weeks of protection within a 12 month period, albeit unpaid, to mothers under the Family and Medical Leave Act (FMLA).[12] There are two conditions, firstly the firm must be large with at least 50 employees and secondly the women must have accrued at least 1250 hours of employment in the past 12 months.

[10] See OECD Family Database.

[11] These states were California, Connecticut, District of Columbia, Maine, Massachusetts, Minnesota, New Jersey, Oregon, Rhode Island, Tennessee, Vermont, Washington and Wisconsin.

[12] The FMLA covered absenses from work for other reasons than maternity, such as the need to care for an ill family member.

In the states offering unpaid leave prior to FMLA, the period of time for which the leave was available varied somewhat. The details of the timing of implementation of the policy and the period of eligibility, detailed in Appendix Table 1, show that some states were more and some less generous than FMLA (for example California offered 6 months leave). Since 1993 some states have made changes to the eligibility rules for FMLA in terms of weeks of leave, the previous 12 month hourly requirement and the firm size[13], although the timing of these changes was outside the sample period in this paper. The most significant departure was in 2004, where three states have attached a pecuniary benefit to the leave, California being the first to do this.[14] Our sample period runs up to 1997, before the introduction of state-level paid maternity leave.

In our data, firm size is not observed and the hours worked by women is recorded for a calendar year rather than the 12 months prior to pregnancy. Our definition of FMLA coverage is initially limited to an on-off treatment. FMLA is set equal to one if children are born on or after the month FMLA was implemented in the state[15], in the period of eligibility (predominantly 12 weeks) and if the mother worked during pregnancy. It is set equal to zero otherwise, where we control for pre-pregnancy employment status of the mother. To take account of the eligibility requirement to work 1250 hours in the year before birth we use the annual data on hours to i) estimate the effects of unpaid maternity leave on the female labour supply mechanism, distinguishing between movements into part- and full-time work (section 5.3) and ii) run a placebo test on mothers working less than 1250 hours pre-birth (section 7.1).

3 Data and Descriptives

The main data comes from the US Panel Study of Income Dynamics (PSID). The PSID comprises a representative sample of households followed since 1968. Members in the household were followed annually until 1997 and then biennially until 2009. As the survey is

[13] See Espinola-Arrendondo and Mondal (2010) for details. Only 4 states expanded FMLA for both private and public sector employees. These were Conneticut, Maine, Oregon and Vermont.

[14] Rossin-Slater et al. (2013) provide an evaluation of the extension in California.

[15] including states which implemented unpaid maternity leave prior to 1993

household based, information is collected on any descendents, creating a cross-generational dataset. Monthly retrospective labour market status, including employment, time out of the labour force and unemployment was available for each year between 1984-1997 and this is our sample period. We constructed a sample of male-female pairs and follow the parents from when the child was first born up until the month the mother either re-enters work or has another child. A state-change to employment or fertility is an absorbing state and no further observations for that mother are included in the sample.

There are 2340 children in the sample and 1560 households. With the time dimension, there are 30664 observations.

Both women and men report whether they were employed for at least part of each month of every month of the previous year. The indicator for exogenous paternal employment is generated from a question asking whether the respondent's job changed since January in the previous year. If it did, the reason for the change is recorded. Similarly to Rhum (1991) and Stevens (1997) the job loss is assumed exogenous if the reason stated was from plant closure, being laid off or fired.[16] A potential problem with identification is that fathers fired from jobs may be non-random in the population and these are included in the involuntary job loss variable. However according to Boisjoly et al. (1994), only 15.7% were fired and the majority were laid off. We create a monthly measure of exogenous job loss (layoff), by combining this information with monthly retrospective data on the individual being unemployed but looking for work. The indicator for monthly layoff is a contemporaneous measure which takes the value of 1 in the month of layoff and 0 during a month of no layoff. In addition a measure for previous layoff takes the value of 1 if the partner was laid off in the 5 years prior to childbirth and 0 otherwise.

To identify the point in time when the mother returns to work, women are defined as being in employment if they report having a job for the entire month, having worked all weeks, or if they missed one or more weeks but not because of a layoff. Data on state level maternity leave legislation was taken from Baum (2003b). In addition, a measure of working during pregnancy takes the value of one if the mother reported working in the labour market between one and nine months before birth.

[16] It is not possible to identify which of these three reasons applied to the individual job change.

Additional data is merged at the state level onto the PSID in order to control for labour market opportunities for women. We collect monthly state level female employment information from the Current Population Survey (CPS) Merged Outgoing Rotation Groups (MORG), available since 1979. Employment rates of the civilian labour force aged 16 or over are calculated for each state, stratifying by the education of the woman. Education levels are defined as having no qualifications, a high school or finally a degree (a 4 year college degree or more).

Table 1 shows the sample statistics. The child-level sample includes families across time from the birth of each child. 7.9% of fathers are exogenously laid off from their jobs after the birth of a child and before the mother changes state and 8.9% are laid off within the first 3 months of the child birth. On average 28% of the observations live in states with some unpaid maternity leave legislation and 6.8% of observations are eligible for FMLA (which means they live in a state with unpaid maternity leave, worked during pregnancy and their child is no older than the eligibility period).[17] Note that of the 8.9% of fathers laid off within the first 3 months, 16.2% of these are treated.

15.6% of fathers have experienced layoff in the past, or 39.0% of fathers conditional upon contemporaneous monthly layoff taking the value of 1. 41.0% of mothers report working during pregnancy. Looking at the measures for parental socio-economic status, 23.5% of fathers and 18.3% of mothers have a degree and the average age of fathers and mothers is 33 and 30 respectively. The average employment rate for women (stratifying by education) is 60.5%.

Finally the table details statistics relating to the age that the mother moves labour state[18], to re-enter employment (12807 mothers with a mean of 26 months), have another child (5181 mothers with a mean of 32 months) or never change their labour state (12676 mothers).

We observe the mother to be at home with their child or to have shifted to a destination labour state - either employment or fertility. The destination labour states are absorbing

[17] Section 5.3 extends the treatment definition to include the hours worked in the previous year using annual data.

[18] to avoid confusion between terminology for the state of residence and destination state (whether the mother is at home, works or has another child), the latter will be referred to as the labour state.

and panel a) of Table 2 reports that 72.78% of mothers enter work, 17.39 are censored and 9.83% have another child. Panel b) of Table 2 shows that in any month 93.7% of mothers are at home with their child, 5.55% are in employment and 0.75% are in a fertility state which means that in that month they have a child.

4 Methodology

The paper aims to identify the extent to which a mothers' labour supply response to her spouse's job loss depends on the availability of unpaid maternity leave benefits, using a discrete time duration model.[19] Specifically, using a difference-in-difference approach the paper asks whether mothers speed up their return to work in response to a paternal employment shock and how maternity leave coverage affects this response.

The choice of mothers to re-enter employment after childbirth is likely to be taken simultaneously with the choice over further fertility. We model three choices of a mother in any time period t - to remain at home with the first child, to have another child or to (re-)enter the labour market conditional upon being at home in period $t-1$ where t refers to the age of the child in months. We use a competing risk model, where there exist two mutually exclusive absorbing destination labour states $j = \{w, f\}$, where w and f denote work and fertility. The hazard function at time t, $h(t)$ is the sum of the hazard for destination to labour state w and f ($h_w(t)$ and $h_f(t)$ respectively). The hazard for exit to labour state w (f) is given by the probability of exit to w (f) in period t, given the individual remained at home with the child up to period $t-1$. We refer to this below as the conditional probability. Let y_{it} denote the labour state of mother i for child age t months. y_{it} takes the value of 0 if mothers do not switch labour state and are at home with their child, 1 if they move to work and 2 if they have another child. The state-specific hazard functions are defined as follows

$$h_j(t) = \Pr(y_{it} = j | y_{it-1} = .. = y_{i0} = 0, x_{it}, layoff_{it}, FMLA_{it}, layoff_{it} * FMLA_{it}); j = \{w, f\} \tag{1}$$

[19] This section follows Dolton and van der Klaauw (1995), Jenkins (1995) and Van den Berg (2001).

where $layoff$ is an indicator taking the value of 1 if the father is exogenously laid off from his job and 0 otherwise. Identification comes from the assumption that layoff is an unexpected shock to paternal employment and we test this assumption in Section 5.5. $FMLA$ is a dummy taking the value of 1 if mother is covered by maternity leave in the month (if they live in a state with unpaid maternity leave, worked during pregnancy and the child is no older than the eligibility period, which was predominantly 12 weeks) and 0 otherwise. The interaction between layoff and FMLA (denoted LF below) is of interest in this paper, it analyses heterogeneity in the layoff effect by eligibility to unpaid maternity leave. x denotes the control variables.

The likelihood function for destination labour state w is then given by

$$\mathcal{L}^{w} = h_{w}(t)\, S(t-1) = \frac{h_{w}(t)}{1 - h_{w}(t) - h_{f}(t)} S(t) \tag{2}$$

where $S(t)$ denotes the survivor function at period t. The likelihood function for destination labour state f $\left(\mathcal{L}^{f}\right)$ is defined similarly.

The data is right censored, as a mother may not return to work or have another child during the sample period, i.e. she may remain at home with her child.[20] The contribution to the likelihood of a censored case $(\mathcal{L}^{c})$ is simply the survivor function

$$\mathcal{L}^{c} = S(t) = \prod_{k=1}^{t} \left[1 - h_{w}(k) - h_{f}(k)\right] \tag{3}$$

The contribution to the likelihood function of an individual is given by

$$\left(\mathcal{L}^{w}\right)^{\delta^{w}} \left(\mathcal{L}^{f}\right)^{\delta^{f}} \left(\mathcal{L}^{c}\right)^{1-\delta^{w}-\delta^{f}} \tag{4}$$

where δ^{j} denotes the exit labour state indicator for $j = \{w, f\}$. Following Allison (1982) we assume a logistic distribution to allow estimation of the duration model by a multinomial logit. In this case the hazard function for exit to labour state w is

[20] The time of censoring, in terms of child age, differs across observations because the timing of births differs across families but all families are observed only up to 1997.

$$h_w(t) = \frac{\exp(\beta_w' \mathbf{X})}{1 + \exp(\beta_w' \mathbf{X}) + \exp\left(\beta_f' \mathbf{X}\right)} \tag{5}$$

where $\mathbf{X}$=($x_{it}, layoff_{it}, D_{it}$). The hazard function for exit to labour state f is defined similarly. Huber-White clustered standard errors are generated at the state level, as part of the variation in access to unpaid maternity leave that is being exploited is across states.

We are interested in the interaction between layoff and eligibility to FMLA.[21] We calculate the marginal effect of layoff on the conditional probability of observing the mother in labour state j at time t for FMLA status equal to zero and one, given she was at home up to period $t-1$. The difference between these conditional predicted probabilities is the effect of unpaid maternity coverage upon the responsiveness of female labour supply to a paternal employment shock.

$$\begin{aligned}
&[\Pr(y_{it} = j | y_{it-1} = .. = y_{i0} = 0, x_{it}, FMLA_{it} = 1, layoff_{it} = 1) - \\
&\Pr(y_{it} = j | y_{it-1} = .. = y_{i0} = 0, x_{it}, FMLA_{it} = 1, layoff_{it} = 0)] - \\
&[\Pr(y_{it} = j | y_{it-1} = .. = y_{i0} = 0, x_{it}, FMLA_{it} = 0, layoff_{it} = 1) - \\
&\Pr(y_{it} = j | y_{it-1} = .. = y_{i0} = 0, x_{it}, FMLA_{it} = 0, layoff_{it} = 0)]
\end{aligned} \tag{6}$$

We evaluate the change in the conditional probability in equation (6) averaged across the value of other covariates. The marginal effect will give information on how a paternal employment shock changes the conditional probability of observing a mother in a particular state, given different maternity leave policies. However, the size of the marginal effect will fall as the unit of time in the sample becomes smaller because the probability of moving states falls. We adopt the method of Dlugosz et al. (2014) and normalize the marginal effect by the predicted probability of being in labour state j for an average individual with no paternal shock. This relative marginal effect (RME) is the percentage change in the conditional probability of being in labour state j when changing from no paternal shock to

[21] We follow Ali and Norton (2003) and do not consider the coefficient on the interaction between layoff and treatment from the multinomial logit as the relevant parameter estimate.

a paternal employment shock.

We focus on the paternal employment shock rather than a maternal employment shock for three reasons. The incidence of maternal layoff is very small (1% of mothers are ever laid off in the sample period compared to 10% of men). Second, the incidence of a maternal shock after pregnancy is linked to the eligibility of unpaid leave, as only mothers who remain in employment can be laid off. Finally the paper exploits a paternal shock at a point in time when mothers are not earning which provides a strong incentive for mothers to return to work. On the other hand, when a mother is laid off the father is likely already in work (in the data there are only 11 cases of both being laid off simultaneously).

The control variables we think are important for maternal labour participation include child year of birth dummies, paternal and maternal age and education and maternal year of birth, family size, ethnicity and the state-level female education-specific employment rate.[22] In addition, we include a dummy for previous paternal layoff to control for the fact that whilst the first layoff may be exogenous, further layoffs may be outcomes of the first (and therefore not unexpected). Finally we include a dummy variable for maternal working pre-pregnancy which indicates eligibility for FMLA.

5 Results

5.1 Graphical Results

Figure 2 presents the Kaplan-Meier empirical hazard function for first re-entry to the labour market, considering only the last birth observed for each women, to eliminate the possibility of future movement out of the labour market due to childbirth. The graph distinguishes by mothers eligible and ineligible for FMLA and the hazard functions are plotted against the time since the father was laid off exogenously. Looking firstly at the mothers ineligible for FMLA, Figure 2 shows that the hazard rate is fairly flat across the timing of the layoff. The pattern is noisier when considering eligible mothers, where there is a dip in the hazard three months prior to layoff which then increases up to the point of layoff.

[22] The latter variable is included to address a worry that paternal layoff is more likely in a downturn and as such the women will find it harder to re-enter the labour market.

To test whether the hazard function before layoff is similar for the two samples (eligible and ineligible) we run a log rank test. The null hypothesis, that there is no difference in the hazard by FMLA eligibility is tested against the alternative hypothesis that there is a difference. Taking 1-12 months before the father was laid off, we cannot reject the null that the hazard function for the eligible and non-eligible are the same.[23] But in the period 0-12 months after the father was laid off we reject the null of the same hazard.[24] This suggests that before the shock, the trends for the treated and the control groups are not statistically significantly different to each other but a divergence occurred after the paternal layoff.

The increase in the hazard leading up to the layoff for FMLA eligible mothers may be noise or alternatively due to the layoff being predictable in the months before the event occurred. There is no reason to think that this would be different depending upon eligibility to FMLA . In any case, as a robustness check in Section 5.5, we redefine the period of layoff to include the prior 3 periods and find no significant effect of the new definition, suggesting that mothers' labour supply does not react to layoff prior to the event. Next a competing risk analysis allows for more flexible treatment of further fertility than restricting the graph to the last observed birth.

5.2 Regression Results

Table 3 reports the marginal effect of a paternal employment shock upon the conditional probability of observing a mother in state j, where $j = \{w, f\}$ or at home with the child. All regressions control for dummies for previous paternal layoff, maternal working during pregnancy, child year of birth dummies, paternal and maternal age at birth and education, maternal year of birth, family size, ethnicity and the state-education specific female employment rate. The coefficients reported in the tables are evaluated at the actual values of other covariates and the average taken across individuals and regressions cluster at the state level.[25] Standard errors are calculated using the delta method.

Table 3 starts by reporting the baseline effects of paternal layoff and the FMLA eligibility.

[23] The chi-square statistic is 2.76 and with 1 degree of freedom the critical value at 1% level is 3.38.
[24] The chi-square test statistic is 4.48.
[25] Note that analysis that clusters at the individual level does not change to the conclusions of the paper.

The incidence of a paternal layoff has no significant effect on the return to work by mothers. Hence on average there is no evidence of insurance against the shock by movement into the labour market. We define eligibility to FMLA to equal one for eligible mothers in states with some maternity coverage and 0 otherwise. Mothers were only eligible for job protection if they worked for 1250 hours in the year prior to birth (which we proxy with an indicator for working during pregnancy) and protection lasts only for the period of eligibility, rounded to the nearest month (for the majority of observations this was 12 weeks which is measured as 3 months in the data). Hence we set the FMLA indicator equal to zero in states with unpaid maternity leave if either of these criteria are not met.[26] Mothers with access to unpaid maternity leave speed up their return to work relative to mothers with no employment protection by 5.6 percentage points (101.60%). This finding may sound counter-intuitive at first if we expect the FMLA to allow mothers to stay at home for longer with their children. However it is consistent with Berger and Waldfogel (2004) who found that "..women with leave coverage are at a 40% higher risk of returning to work post-birth than mothers without leave coverage" (p. 345).

Of interest to this paper is the difference in the conditional probability from a paternal layoff, by FMLA eligibility. The interaction between FMLA and paternal layoff defined in equation 6 allows us to analyse how the responsiveness of female labour supply after a paternal employment shock varies with maternity leave coverage. The coverage offered new mothers employment protection around the timing of the birth and we hypothesize that it would facilitate the insurance mechanism of female labour supply by reducing search costs involved with returning to work. Column (1) of Table 3 reports a marginal effect of paternal layoff is 2.7 percentage points (49.08%) higher for mothers with employment protection during maternity. If there were no search frictions, this coefficient would be insignificant. For mothers with unpaid maternity protection, paternal layoff raises the conditional probability of returning to work. We see a simultaneous decrease in the probability of mothers being at home with their child. There is a positive but insignificant interaction effect on further fertility choices.

[26] Note that the subsequent section takes into account the further eligibility requirement that the mother must have worked 1250 hours in the year before childbirth, by exploiting data on retrospective hours worked.

Web Appendix Tables W1 reports the marginal effects of the full regression results. Of the variables statistically significantly different to zero, working during pregnancy and maternal education both raise (lower) the conditional probability of working (staying at home) whilst parental age, mother year of birth and family size have the opposite effect. The child year of birth dummies (omitted) show maternal labour supply to be increasing significantly across the period of observation.

5.3 Part-Time and Full-Time Employment

In a model of female labour supply, Francesconi (2002) notes the need to distinguish between full- and part-time employment, as each exhibits a different wage-experience profile and are differentially substitutable with leisure. In the data, 37% and 30% of women work full- and part-time respectively. Table 4 estimates the role of maternity leave coverage in driving female labour supply as an insurance mechanism, looking at women returning to part-time work (column 1) and full-time work (column 2). In the PSID data, part- and full-time work history is available only on an annual basis rather than monthly. In order to extend the analysis to distinguish between part- and full-time work we must assume that mothers do not change between the two labour states within a year.

The baseline effect of paternal layoff is again insignificant. The baseline FMLA effect significantly shifts mothers from staying at home towards part- and full-time employment. Mothers with access to unpaid maternity leave have a conditional probability of being in part- and full-time work 2.6 and 2.8 percentage points (or 89.85% and 109.35%) higher than mothers with no leave, respectively.

Looking at the difference in the effect of layoff by FMLA eligibility, the effect of a paternal shock raises the conditional probability of returning to part-time work by 1.4 percentage points more for FMLA eligible mothers, but the effect is insignificant and it is significantly different to the comparable result of Table 3 column 1 at the 10% level.[27] On the other hand, the difference in the effect of layoff on the conditional probability of returning to full-time work by eligibility to unpaid leave is statistically different to zero and instatistically different

[27] The t-statistic for the hypothesis of identical coefficients is 1.8.

to the estimate in Table 3.[28] Eligible mothers have a higher conditional probability by 1.2 percentage points, or 45.51%. We would expect a larger shift to full-time work and not part-time work, as access to FMLA required at least 1250 hours worked in the previous 12 months so part-time workers are less likely to be eligible. Note that mothers returning to part-time work may have been eligible for FMLA, but reduced their hours after having a child.

In summary, we do see a significant speed up of the return to full-time work (and an insignificant speed up to part-time work) for mothers with access to unpaid maternity leave.

5.4 Search Frictions

This paper suggests that the mechanism for an insurance role of unpaid maternity leave is through search frictions. Unpaid leave allows a mother to re-enter the labour market without exerting job search effort. If there were no search frictions, a mother ineligible for unpaid leave could quickly and costlessly find a new job when desired. This section explores possible heterogeneity in the insurance role of FMLA, by varying levels of three variables that proxy for search frictions. These are i) the business cycle, ii) the local labour market conditions and iii) the level of maternal education.

Menzio and Shi (2011) find individual movements from unemployment to employment in the US are positively correlated with the business cycle. We define a business cycle using log real GDP for each quarter, between 1947-2015 from the US Bureau of Economic Analysis.[29] To remove quarterly and annual trends in log real GDP, it is regressed on quarter and year dummies and a residual predicted to proxy the quarterly business cycle fluctuations. A peak in the business cycle is defined as fluctuations in the 75th percentile and above. Columns 1-2 of Table W2 shows a smaller magnitude for the interaction between FMLA and layoff during a peak in the business cycle, as expected. The marginal effect of paternal layoff is 1.6 percentage points higher for FMLA eligible mothers (30.21%) compared to 2.2 percentage points (39.72%) during a non-peak business cycle, although the difference is not statistically

[28] The t-statistic is 1.6.

[29] This includes data outside of the sample period to improve the accuracy of the estimated business cycle fluctuations.

significant.

Motivated by a similar logic, columns 3-4 analyse heterogeneity in the insurance role of FMLA by distinguishing between areas of high and low female unemployment, in the local labour market. We expect search frictions to be relatively stronger in areas of high unemployment. However the results show very similar marginal effects of FMLA by layoff status for the high and low unemployment areas.

Finally, high educated workers face greater search frictions in the labour market. They have a relatively higher reservation wage and the process of acquiring specific skills reduces the number of potential employers. This heterogeneity in search frictions is reflected in the results in columns 5-6 of Table W2 which show that the magnitude of the interaction between FMLA and layoff is higher for high educated mothers. Again difference between the two groups is not statistically significant.

The sample of households with a paternal shock and unpaid maternity leave is relatively small and therefore there is little power to detect significant heterogeneity. Despite this looking at the magnitudes alone, the results do show that the insurance role of unpaid maternity leave is stronger when search frictions are greater.

5.5 Endogeneity of Paternal Layoff

In this section we consider two potential sources of endogeneity of paternal layoff. Firstly, the layoff may be an outcome of previous separation from the labour market and therefore predictable and secondly the employees may predict the layoff and therefore respond before the event is realized.

The potential endogeneity of the paternal employment shock may cause a bias if, for example the father has been laid off in the past. Stevens (1997) notes that being laid off from work leads to future job uncertainty which can take the form of further layoffs. The layoff variable is interpreted in the paper as a shock to employment status, but for fathers with past lay-off experience it may be an outcome of a past event rather than an exogenous change and consequently predictable by the mother. Table 5 drops from analysis any households where the father has experienced a layoff in five years leading up to the birth thereby repeating analysis on a sample with cleaner identification.

Before discussing the results it is useful to think about what we would expect to find. If the estimated insurance effect of FMLA is driven by households with previous experience of paternal layoff then excluding them from the sample should see a drop in the magnitude. We find instead no statistically significant difference in the estimates of Table 3 and 5 in the interaction term[30], suggesting that the earlier results were not biased by endogenous paternal layoff.

This leads to the next question of whether the first layoff is really exogenous. It could be argued that the first layoff experienced is not a shock, if "bad" fathers who are laid-off live with mothers who return to work quickly after childbirth. However this is true in all households which experience a first layoff shock, irrespective of unpaid maternity leave eligibility. But as we find very different effects of the shock by eligibility, it suggests that this bias cannot drive our results.

We consider the second source of endogeneity in response to the evidence in Figure 2, that the hazard for return to work after childbirth increases in the months leading up to layoff for mothers eligible for FMLA. Web Appendix Table W3 redefines the definition of layoff to include the three months prior to the layoff. The intuition is that employees knew in advance that the firm was due to close and started to react to the event before it actually occurred. The consequence of including the additional months is that the interaction term is no longer significant for each of the three states and the coefficients are statistically different to the respective coefficients in Table 3[31], suggesting that in fact mothers do not respond prior to the event itself, by returning to work.

6 Maternity Leave as Insurance

The paper hypothesizes that unpaid maternity leave offers insurance possibilities to households by allowing the mother to return to work in the event of a paternal shock. This section provides further tests for the hypothesis by firstly analysing the response of the timing of return to work to a paternal shock, differentiating between those eligible or not for *paid*

[30] The t-statistic for the hypothesis test of equality coefficients is 0.04.

[31] the t-statistics are 3.07, 3.43, 6.64 for states work, home and fertility respectively.

maternity leave. The paid leave we analyse is not a full reimbursement of wages but has a value similar to unemployment benefits. Nonetheless during a period of paid leave, the increased income from the mother returning to work will be lower than for mothers ineligible for paid leave and therefore we expect to find a smaller insurance role of paid leave compared to unpaid leave. Secondly the section examines the food consumption response to a paternal shock by households eligible and non-eligible for FMLA. If the unpaid maternity leave were acting as an insurance mechanism then we would expect to see smoothing of food consumption against the paternal layoff shock, for those eligible for FMLA.

6.1 Paid Maternity Leave

The analysis so far has considered only the role of unpaid maternity leave in providing insurance for households. However Lalive et al. (2014) highlight the importance of an additional form of maternity leave which is cash benefits and find that time to care for children just after birth is greatest for systems which combine the both paid and unpaid leave. Five states in the US offer a form of paid maternity leave, through the Temporary Disability Insurance (TDI).[32] The states allow payment of a benefit for the period when mothers are "disabled" after childbirth, of similar value to unemployment benefit (see Espinola-Arrendondo and Mondal, 2010 for details). This period is generally thought to be around six weeks, although will vary depending upon the health of the mother. We supplement the analysis by evaluating the effect of paid maternity leave upon the responsiveness of female labour supply to a paternal layoff.

We hypothesize that whilst unpaid maternity leave insures households against shocks by speeding up the mothers' return to work, paid leave offers a benefit payment which itself smooths the effect of a paternal shock, allowing the mother to delay her return to work. To test this hypothesis, we define the variable TDI to take the value of 1 in TDI states, whose child is no older than 2 months old and 0 otherwise.

The results are displayed in Table 6. To avoid conflating the effect of the TDI and FMLA policies, mothers eligible for FMLA are dropped from analysis, leaving a sample of

[32] These states are California, Hawaii, New Jersey, New York and Rhode Island

29587. Similarly to the specification above, Table 6 reports the marginal effects where the other covariates are evaluated at their values and an average of the probabilities taken across individuals. The regressions cluster at the state level.

Similarly to above, the coefficient layoff is small insignificant and the coefficient on the TDI variable shows that mothers with access to TDI speed up their return to work. The next rows reports the coefficient of interest, the difference in the marginal effect of paternal employment shock by TDI eligibility. The parameter is equivalent to equation 6, substituting FMLA eligibility for TDI eligibility. The marginal effect of paternal layoff is 0.8 percentage points (10.03%) higher in households with potential TDI eligibility but the coefficient is insignificant. In addition the coefficient is statistically different to the comparable estimate in Table 3.[33] In this case, we find no significant heterogeneity in the effect of layoff by access to TDI.

It is true that the analysis is crude and requires better data on exactly who received the benefit and for how long. For example we can identify whether a mother lived in a state which offered TDI but we cannot tell for how many weeks the benefit was received, if at all. However, the preliminary analysis suggests that whilst on a period of paid maternity leave, mothers in receipt of a paternal layoff do not speed up their return to work. We discuss further in the conclusion that, as the early return to work has consequences for child development, this finding suggests that mothers should have access to paid maternity leave in early weeks in order to limit the impact of a paternal layoff whilst protecting child human capital.

6.2 Food Consumption

There is a large literature estimating the extent to which households insurance themselves against income shocks by examining the response of food consumption. A natural question to ask is whether we observe insurance effects of FMLA on food consumption. In the PSID food expenditure is reported annually from as early as 1968 however prior to 1975 the measure of consumption is somewhat incomplete as it refers only to annual expenditure on food at

[33] The t-statistic for the test of equality of coefficients is 4.00.

home and we restrict analysis therefore to 1975 onwards. Useful for our analysis is that we can distinguish food expenditure using food stamps. In the event of a paternal layoff, the government in the US offers a form of insurance in the form of food stamps. Consequently the measure for total food expenditure may respond less to the paternal employment shock than the measure excluding the food stamps. We define three measures of annual food expenditure - total, food consumption minus food stamps and food consumption in the home. The mean value of each measure is \$4070, \$3760 and \$3079 respectively, reported in 2000 prices.

As the food consumption data is measured annually, we construct a measure of layoff which takes the value of one if the partner was laid off in a particular year and 0 otherwise. We implement the following difference-in-difference estimation using OLS

$$\ln c_{it} = \beta_0 + \beta_1 layoff_{it} + \beta_2 FMLA_{it} + \beta_3 layoff_{it} * FMLA_{it} + \beta_4 z_{it} + \varepsilon_{it}$$

where $\ln c_{it}$ denotes log annual consumption for household i in year t, $layoff$ and $FMLA$ are defined as above and z denotes a set of covariates including paternal and maternal age and education and maternal year of birth, family size, ethnicity a dummy variable for maternal working pre-pregnancy and for previous paternal layoff aswell as year and child age dummy variables.

The results are reported in Table 7. The sample includes more years of data and is larger than in the original analysis.[34] There are three columns of results, one for each consumption measure. The sample size is larger than in the main tables, owing to additional years of data on food consumption included in the analysis. The effect of a paternal layoff is to reduce annual food consumption by 3.7%, 5.5% and 5.5% for total food, household food and food excluding stamps respectively where all three coefficients are statistically significant. On the other hand, no statistically significant effect of FMLA is found on food consumption, as is to be expected as maternity leave eligibility should not alone drive consumption. But importantly for the analysis, the coefficient β_3 suggests that the negative effect of layoff is removed for mothers eligible for FMLA. In fact the coefficient is large at between 8-13% which cancels out the effect of layoff. Notably, there is some evidence that the insurance role of

[34] Repeating the analysis on the sample from Table 3 yields the same conclusions.

FMLA is mediated by food stamps, as the interaction coefficient is slightly smaller in the first column (which includes food stamps) than in the second column of results. Taken together, the finding of consumption smoothing against the paternal shock for mothers covered by unpaid maternity leave provides further evidence of the insurance role of unpaid maternity leave.[35]

7 Placebo Tests

7.1 Mothers Working Part-Time in the Year Before Birth.

As mentioned in Section 2, an eligibility criteria for unpaid leave was the accrual of 1250 hours of work in the 12 months leading up to the birth. We would expect no interaction effect of FMLA with the paternal layoff for mothers working fewer hours, which makes for an interesting placebo test.

Hours worked is available annually rather than monthly and we run a placebo test by repeating analysis of Table 3 on the sample of mothers working less than 1250 hours in the year before birth. For space considerations, the placebo tests in Table 8 report only the marginal effect of the interaction between FMLA and layoff on the conditional probability of observing the mother in each of the three states (work, home and fertility). Panel 1 distinguishes between part-and full- time work in the year before birth, with panel 1a reporting the placebo effects for ineligible mothers working less than 1250 hours and panel 1b for the mothers working at least 1250 hours. Panel 1a shows no significant difference in the marginal effect of paternal layoff by FMLA status. Moreover the estimate is statistically different to the estimate in Table 3.[36] Hence we find no effect on the placebo group of ineligible mothers. For eligible mothers in panel 1b the difference increases to 3.6 percentage points (63.13%) but again becomes insignificant. In this case, there is no statistical difference between the estimate in panel 1b and Table 3.[37]

[35] Similar conclusions are drawn from regressions restricting the sample years as in the main analysis. The coefficient (standard errors) for β_3 are 0.076(0.037), 0.080(0.044).0.068(0.049) for column 1, 2 and 3 respectively.

[36] The t-statistic for the test of equality of coefficients is 2.89.

[37] The t-statistic is 0.17.

7.2 Setting a Randomly Chosen Implementation date for FMLA

We may worry that the insurance effect of FMLA estimated in Table 3 is biased and picking up other unobservable attributes of treated areas and treated mothers. To investigate this, a placebo experiment creates six artificial dates for the implementation of FMLA up to one year before the federal policy change in August 1993. We exclude the six months immediately before the policy change as many of the mothers with placebo treatment status equal to one are also treated in reality. If we have correctly specified the model, we should find no significant insurance effect from FMLA implemented at these false policy implementation date.

What is reported in panel 2 of Table 8 is the difference in the marginal effect of layoff by eligibility for unpaid maternity leave. The table shows that in all cases the coefficients are statistically insignificantly different to zero and the magnitude of estimates are lower than in our main regression. Furthermore in almost all cases the coefficients from columns 1 and 2 are statistically different to the respective columns in Table 3 – suggesting that whilst the real policy led to an insurance effect, the placebo policies did not. Note that in column 3 of the placebo test table, some coefficients are not statistically different to column 3 of table 3, if in both cases the marginal effects are insignificantly different to zero. Where the coefficients are significant in the main paper, they are statistically different to those in the placebo tests.[38]

8 Conclusion

If there were complete markets, households would fully insure themselves against shocks to employment. This paper has shown that with maternity leave restrictions mothers are less able to effectively use their labour supply to smooth shocks to paternal employment. In particular, in the absence of employment protection during maternity, the conditional probability of a mother returning to work in response to a paternal employment shock is 49.08% or 2.7 percentage points lower than for mothers eligible for the unpaid maternity

[38] Selecting different false implementation dates led to similar results, for example 6 months and 1 year before and after August 1993. Results are available from the author on request.

leave. The insurance role of unpaid maternity leave was supported by additional findings which suggested that in the 5 states which offered *paid* maternity leave, the female labour supply response to a layoff was smaller than for mothers with unpaid leave, and statistically insignificant and by evidence that the negative effect of a paternal layoff on household food consumption was smoothed out for mothers covered by maternity leave. Whilst to date a large literature has evaluated the benefits of unpaid maternity leave in terms of maternal and child outcomes, this is the first paper to add a third benefit of the insurance role.

To give the effect some magnitude, the mean local employment rate of females in the sample was 62% and the standard deviation 1%, suggesting we estimate large shifts into work. Figure 1b showed the female participation rate to drop from 70% pre-pregnancy to around 55% thereafter, a change of 15 percentage points. So the additional insurance response of female labour supply to a paternal shock in areas which protect employment for a spell around childbirth is 20% the size of this change. It seems that given the opportunity, households are keen to take advantage of the ability to self-insure against shocks.

If households are not adequately able to insure themselves there will be welfare consequences both to the adults in the household in terms of consumption but also to the children. Children living in households that experience negative shocks tend to accumulate lower levels of education, have lower earnings are more likely to drop out from high school (Carneiro et al. 2015b). However, whilst it is positive that the unpaid maternity leave improved insurance possibilities for families, a speed up of the return to work within the first 12 weeks of birth could be harmful for the child. Of the vast literature on the effects of maternal labour supply upon child human capital outcomes[39], only a small number of papers focus on mothers returning to work within the first 12 weeks of life. Berger et al. (2005) find a speed up of the return to work by mothers after childbirth to be associated with a lower incidence of breast-feeding and immunization, and a higher prevalence of child externalizing behavioural problems. Baum (2003a) found that the increase in income associated with a mother working within the first 3 months did not fully offset the negative effect on vocabulary scores. Interestingly, among the sample of mothers who did eventually return to work, Baum found a negative effect of hours worked in the first quarter, suggesting that if mothers were able

[39] see Blau and Currie (2003) for a review

to take more time off after childbirth, it would result in an improvement in child outcomes. Finally, Carneiro et al. (2015a) found that a policy in Norway to extend maternity leave from 12 weeks of unpaid to 4 months of paid leave raised time that mothers spent with their children and improved long-term human capital outcomes for the children. Brooks-Gunn et al. (2002) however found an insignificant effect of mothers returning to work by the third month on cognitive ability measures. The general conclusion of these papers indicate that early return to work is harmful to children. Combining with the finding of a statistically insignificant effect of paid maternity leave on the timing of return to work, the evidence suggests that, an extension of the FMLA to offer paid maternity leave may protect not just income levels but also child human capital against household shocks.

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Figure 1a

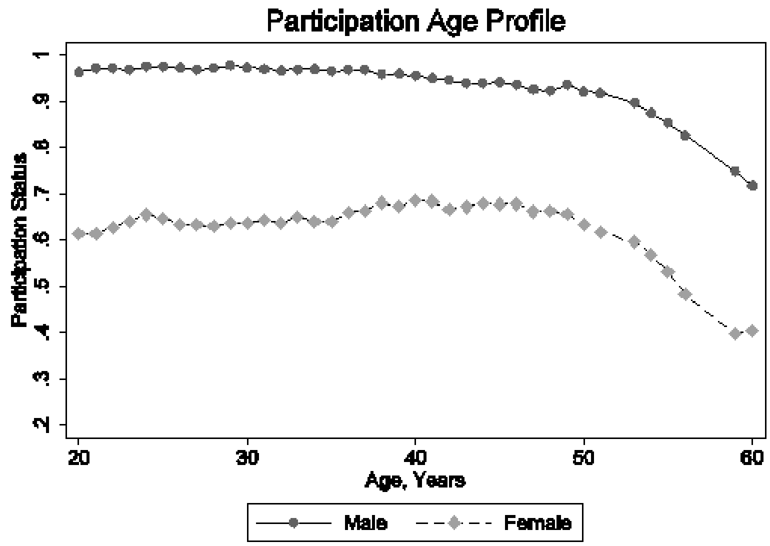


Figure 1b

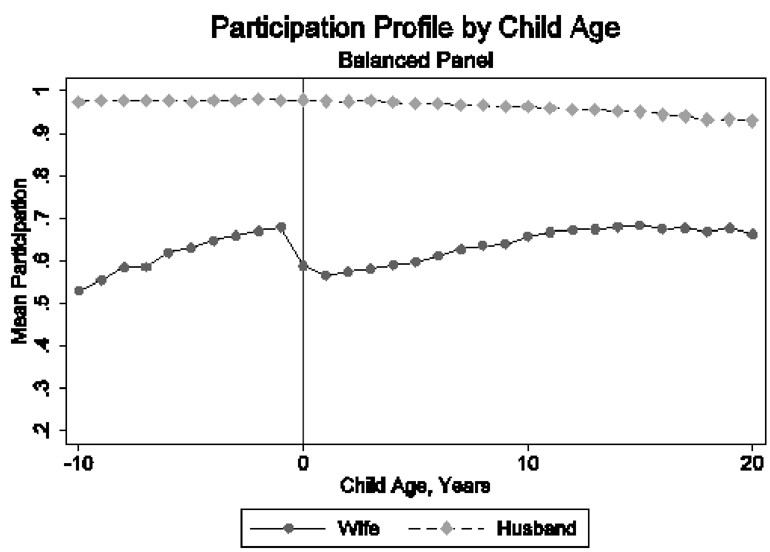
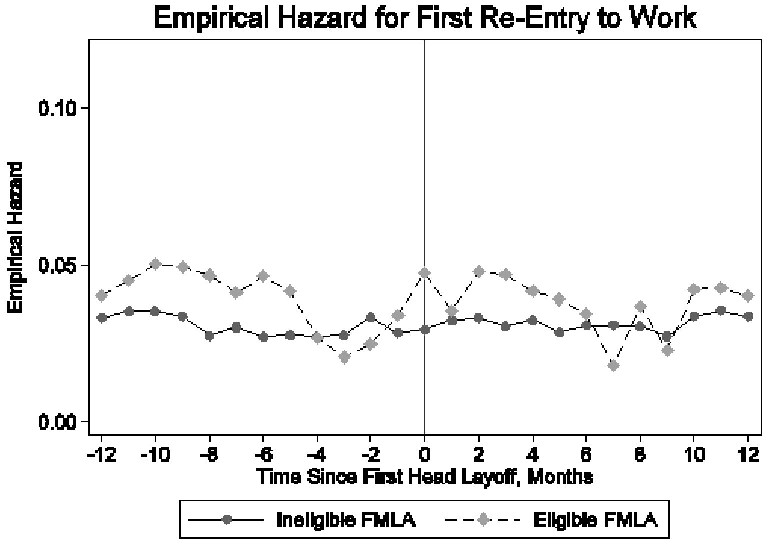


Figure 2: First Re-Entry to the labour Market.



Sample includes last births. Hazard rate calculated as # spells ending during t / # spells not ended in t-1.

Table 1: Sample Statistics

Variable	Sample	Mean	Standard Deviation
Paternal Monthly Layoff	30664	0.079	0.269
Paternal Layoff during first 3 Months of Birth	6,346	0.089	0.284
Maternity Leave Legislation Indicator	30664	0.280	0.449
FMLA Eligibility Indicator	30664	0.068	0.251
Previous Layoff Indicator	30664	0.156	0.363
Maternal Work During-Pregnancy Indicator	30664	0.410	0.492
Father Degree Status	30664	0.235	0.424
Mother Degree Status	30664	0.183	0.387
Father Age	30664	33.156	6.881
Mother Age	30664	30.299	5.861
Local Education-Specific Female Employment Rate	30664	0.605	0.096
Family Size	30664	3.059	1.649
Age return to work	12807	26.06	22.81
Age mother has next child	5181	32.24	17.30
Age of final observation if censored	12676	57.68	29.54

Local education specific female employment rate matches the state specific female employment rate by education categories no qualifications, high school or degree +.

Table 2: Duration Statistics: Monthly Maternal State

State	Sample	Proportion
a) Destination for each Child		
Work	1,703	72.78
Censored	407	17.39
Fertility	230	9.83
Total	2,340	100
b) In a particular month		
At home with Child	28731	93.70
Work	1703	5.55
Fertility	230	0.75
Total	30664	100

Table 3: Competing Risk Estimation of Labour Market Entry by Maternity Leave Policy

	(1)	(2)	(3)
	Return to Work	At Home with Child	Further Fertility
Layoff	0.005	-0.005	0.001
	(0.006)	(0.006)	(0.002)
RME	8.77%	-0.57%	6.74%
FMLA	0.056***	-0.058***	0.003
	(0.006)	(0.008)	(0.003)
RME	101.60%	-6.22%	36.17%
Layoff * FMLA[1]	0.027**	-0.036***	0.009
	(0.012)	(0.015)	(0.007)
RME	49.08%	-3.85%	123.18%
N=30664			

Marginal effects reported. [1] Difference in effect of paternal shock by FMLA eligibility defined in equation (6). FMLA = 0 if no maternity coverage or maternity coverage but mother did not work pre-pregnancy or child no older than eligibility criteria. Controls include paternal and maternal age and education, maternal year of birth, family size, ethnicity, previous paternal layoff, maternal working pre-pregnancy and state education-specific female employment rate, child year of birth dummies. RME is relative marginal effect: relative to probability of being in state for average individual with layoff=0. Standard errors clustered by state and computed by Delta method. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 4: Distinguishing between Part-Time and Full-Time Status

	(1)	(2)	(3)	(4)
	Return to PT Work	Return to FT Work	At Home with Child	Further Fertility
Layoff	0.001	0.004	-0.005	0.000
	(0.004)	(0.003)	(0.005)	(0.002)
RME	2.81%	14.18%	-0.52%	6.23%
FMLA	0.026***	0.028***	-0.059***	0.004
	(0.004)	(0.003)	(0.006)	(0.003)
RME	89.85%	109.35%	-6.30%	58.90%
Layoff * FMLA[1]	0.014	0.012*	-0.035***	0.009
	(0.010)	(0.007)	(0.014)	(0.006)
RME	48.19%	45.51%	-3.72%	121.48%
N=30664				

Table 5: Selecting Households with No Previous Layoff Status of Father

	(1)	(2)	(3)
	Return to Work	At Home with Child	Further Fertility
Layoff	0.005	-0.005	0.001
	(0.006)	(0.006)	(0.002)
RME	8.56%	-0.56%	6.72%
FMLA	0.055***	-0.058***	0.003
	(0.006)	(0.005)	(0.003)
RME	100.49%	-6.18%	36.20%
Layoff * FMLA[1]	0.032*	-0.042*	0.010
	(0.018)	(0.022)	(0.009)
RME	58.28%	-4.47%	131.35%
N=25869			

Marginal effects reported. [1] Difference in effect of paternal shock by FMLA eligibility defined in equation (6). Controls same as Table 3. RME is relative marginal effect relative to probability of being in state for average individual with layoff=0. FMLA = 0 if no maternity coverage or maternity coverage but mother did not work pre-pregnancy or child no older than eligibility criteria. Standard errors clustered by state and computed by Delta method. *** $p<0.01$, ** $p<0.05$, * $p<0.1$.

Table 6: Paid Maternity Leave Through Temporary Disability Insurance

	(1)	(2)	(3)
	Return to Work	At Home with Child	Further Fertility
Layoff	0.004	-0.004	0.001
	(0.006)	(0.006)	(0.002)
RME	6.83%	-0.46%	8.32%
TDI	0.030***	-0.031***	-0.002
	(0.004)	(0.005)	(0.005)
	55.09%	-3.62%	-32.01%
Layoff * TDI[1]	0.005	0.071***	-0.077***
	(0.013)	(0.017)	(0.007)
RME	10.03%	7.61%	1026.97%
N=29587			

Table 7: Insurance of Food Consumption

	(1)	(2)	(3)
	Log Food Consumption (total)	Log Food Consumption (excluding Food Stamps)	Log Food Consumption (within house only)
Layoff	-0.037**	-0.055**	-0.055**
	(0.017)	(0.023)	(0.022)
FMLA	0.004	0.000	-0.006
	(0.017)	(0.017)	(0.017)
Layoff * FMLA	0.081*	0.109**	0.127***
	(0.041)	(0.046)	(0.042)
N	68,554	68,554	68,554

Marginal effects reported. [1] Difference in effect of paternal shock by FMLA and TDI eligibility defined in equation (6). Controls same as Table 3. RME is relative marginal effect relative to probability of being in state for average individual with layoff=0. TDI=1 in TDI states for child no older than 2 months and 0 otherwise. FMLA = 0 if no maternity coverage or maternity coverage but mother did not work pre-pregnancy or child no older than eligibility criteria. Standard errors clustered by state and computed by Delta method. *** p<0.01, ** p<0.05, * p<0.1.

Table 8: Placebo Tests

The marginal effect of paternal employment shock

	(1)	(2)	(3)
	Return to Work	At Home with Child	Further Fertility
1) Heterogeneity by hours worked in year before birth			
a) Worked less than 1250 hours in year before birth			
Layoff * FMLA	0.007	0.065***	-0.073***
	(0.015)	(0.016)	(0.009)
RME	12.17%	6.63%	968.88%
b) Worked at least 1250 hours in year before birth			
Layoff * FMLA	0.036	0.089	-0.125***
	(0.056)	(0.062)	(0.015)
RME	63.13%	9.47%	1753.38%
N=28030			
2) Setting false policy implementation dates			
a) August 1992	0.015	-0.022	0.007
	(0.015)	(0.017)	(0.006)
t-statistic for test of same coefficient	2.052	-2.253	1.818
b) September 1992	0.018	-0.026	0.007
	(0.015)	(0.018)	(0.006)
t-statistic for test of same coefficient	1.539	-5.263	1.818
c) October 1992	0.022	-0.030	0.009
	(0.016)	(0.019)	(0.006)
t-statistic for test of same coefficient	2.791	1.038	0.000
d) November 1992 [9]	0.019	-0.027	0.008
	(0.016)	(0.019)	(0.006)
t-statistic for test of same coefficient	4.465	1.558	0.909
e) December 1992	0.019	-0.027	0.008
	(0.016)	(0.018)	(0.006)
t-statistic for test of same coefficient	4.465	-4.737	0.909
f) January 1993 [5]	0.019	-0.027	0.008
	(0.016)	(0.019)	(0.006)
t-statistic for test of same coefficient	4.465	1.558	0.909
N==30664			

[1] Difference in effect of paternal shock by FMLA eligibility defined in equation (6). Controls identical to Table 3. Standard errors clustered by state and computed by Delta method. *** p<0.01, ** p<0.05, * p<0.1.

9 Appendix

Appendix Table 1: States Early to Implement Unpaid Maternity Leave

State	Date Implementation	Period of Eligibility
California	January 1992	6 months
Connecticut	July 1990	16 weeks
District of Columbia	April 1991	12 weeks
Maine	April 1988	12 weeks
Massachusetts	October 1972	12 weeks
Minnesota	July 1987	6 weeks
New Jersey	April 1990	12 weeks
Oregon	January 1988	12 weeks
Rhode Island	July 1987	13 weeks
Tennessee	January 1988	4 months
Vermont	July 1992	12 weeks
Washington	September 1989	12 weeks
Wisconsin	April 1988	6 weeks

Details for states which implemented unpaid maternity leave prior to 1993.