# The Long-Run Effects of Wage Replacement and Job Protection: Evidence from Two Maternity Leave Reforms in Great Britain

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**Abstract**

This paper examines the effects of maternity leave coverage on women's employment and career trajectories in Great Britain using data from the British Household Panel Survey. Using a difference-in-differences identification strategy and two changes to the national maternity leave policy, I distinguish between the effects of expanding access to wage replacement benefits and the additional effects of providing job protection benefits. Access to paid maternity leave increases the probability of returning to work after childbirth in the short run, but has no effect on long-run employment. Expanding the amount of job protection available to new mothers results in substantial increases in maternal employment rates and job tenure more than five years later. However, job-protected leave expansions lead to fewer women holding management positions and other jobs with the potential for promotion. Although these maternity leave policies have large employment effects on the extensive margin, there is little evidence of effects on average earnings.

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# 1 Introduction

Women face substantial labor market penalties for having children in many developed countries (Gornick et al., 1998). Mothers earn consistently lower wages, work fewer hours, and are substantially less likely to be employed after having a child compared to before, and these effects persist even 10-20 years after a birth (Kleven et al., 2016). One of the reasons for these long-term motherhood penalties is that mothers tend to experience longer and more frequent employment gaps or prolonged absences in their work histories than do men or childless women. These work interruptions are thought to have large, permanent effects on productivity and firm-specific human capital, as well as on future job prospects and wage trajectories (Mincer and Polachek, 1974; Blau and Kahn, 2000; Spivey, 2005; Lalive et al., 2013). Some policy makers argue that maternity leave may be a way to reduce some of the negative work outcomes associated with having children (Misra et al., 2011). Yet the existing literature on maternity leave shows mixed evidence in support of this claim (Baker and Milligan, 2008; Lalive and Zweimüller, 2009; Rossin-Slater et al., 2013; Schönberg and Ludsteck, 2014; Waldfogel, 1999).

One reason for these mixed findings is that most papers bundle together two important components of maternity leave benefits: wage replacement and job protection. Given the lack of consensus about the effects of maternity leave on employment outcomes as well as the substantial variation in policy design across countries, it is important to understand how the effects of wage replacement and job protection benefits may differ. *Job-protected* paid leave in particular is often proposed as a way to reduce the negative effects of work interruptions by improving employment continuity and increasing firm tenure. However, little is known empirically about the effects of job-protected paid maternity leave benefits compared to wage replacement alone.

This paper examines the differential effects of these two dimensions of maternity leave benefits on women's employment outcomes and career trajectories in Great Britain using data from the British Household Panel Survey (BHPS). Using a difference-in-differences identification strategy and two changes to the national maternity leave policy, I distinguish between the effects of expanding access to wage replacement benefits and the additional effects of providing job protection benefits. In contrast to the existing literature which does not separate these two effects,

I find that having access to paid maternity leave increases the probability of returning to work after childbirth in the short run, but has no effect on employment more than three years after giving birth. On the other hand, expanding the amount of job protection available to new mothers results in substantial increases in maternal employment rates more than five years later. Mothers exposed to a maternity leave policy with increased job protection benefits are 8 percent more likely to be employed five years after the birth of their child compared to mothers with access to wage replacement benefits alone.

Why do wage replacement and job protection benefits have different impacts on long-term labor market outcomes? Using a simple two period labor supply model, I provide some intuition into how these two components of maternity leave policies each affect employment decisions in the period a child is born as well as in a future period. I show that both wage replacement and job protection should increase leave-taking in the period a child is born. But they differentially impact long-term employment rates. If individuals face a wage penalty for taking time out of the labor market, then increasing the wage replacement rate will decrease the share of mothers who are employed in the future. Job protection lowers the search costs associated with re-entry into the labor market and preserves any firm-specific human capital the mother has accumulated. By eliminating or reducing the wage penalty associated with work absences, job protection will increase the long-run employment rate compared to a maternity leave policy with wage replacement alone. This prediction is consistent with my empirical findings, and suggests that job protection benefits are a crucial part of any maternity leave policy intended to increase female labor force participation rates.

If job protection increases firm tenure and there is a positive return to firm-specific experience, it can also potentially increase mothers' long-term earnings and promotion opportunities. However, by increasing the amount of leave that mothers take, job-protected paid maternity leave could also have negative effects on these measures. Extended work absences and non-employment spells negatively affect wage profiles and the probability of future employment (Spivey, 2005). Recent work suggests that family-friendly policies aimed at mothers can reinforce traditional gender roles between mothers and fathers, further reducing the amount time mothers spend in market work (Bergmann, 2008; Gornick and Meyers, 2003, 2008). There is also

considerable concern that maternity leave policies that include long periods of job protection may make it more costly to hire women of childbearing age, causing employers to lower female wages or hire fewer women (Hegewisch and Gornick, 2011; Blau and Kahn, 2016). These negative effects may be especially salient for workers in high-skill careers, where early-career productivity is especially important and absent workers are more difficult to replace (Antecol et al., 2016; O'Flaherty and Siow, 1995; Light and Manuelita, 1995). The motherhood wage gap is much larger for highly educated women than for those with low levels of education, which is consistent with the idea that work absences are more costly for high-skill workers (Anderson et al., 2002). High-skill mothers are also less likely than men to become managers or be promoted, even conditional on measures of observed productivity (Kleven et al., 2016; Gayle et al., 2012; Addison et al., 2014). This suggests that employers may view workers who take leave as less productive or less dedicated to their job even several years later.

An important advantage of the BHPS is that I can analyze firm-specific employment flows and job changes to provide novel evidence on the effect of maternity leave on these outcomes. I find that the availability of job protection raises average job tenure measured five years after the birth of a child by over 13 months and increases the probability of working for the pre-birth employer by 30 percent. However, maternity leave may not level the playing field in terms of the types of jobs that women hold. In Great Britain, job-protected paid leave leads to fewer highly educated women holding management positions and other jobs with the potential for promotion. Although these maternity leave policies have large employment effects on the extensive margin, there is little evidence that they reduce the motherhood penalty in earnings, conditional on employment.

This work makes several important contributions to the existing literature. First, by separately identifying the effects of expanding access to wage replacement benefits from the additional effects of providing job protection, I show that job protection is important in order to increase mothers' labor force attachment in the long run. Second, I provide evidence that access to job-protected paid maternity leave can actually exacerbate gender inequality among highly educated workers in terms of the probability of being promoted or holding management positions.

Finally, understanding the effects of different types of maternity leave benefits can help shape optimal policy in the United States and other countries. Although the U.S. is currently the only developed nation without a paid maternity leave policy, it is becoming increasingly plausible that one could be adopted in the next few years. The Family and Medical Insurance Leave (FAMILY) Act currently proposed in Congress would provide all workers with 12 weeks of partially paid parental leave.[1] While there is no explicit provision for job protection in this bill, my results suggest that it may be an important consideration. The only current source of job-protected leave in the U.S. comes from the 1993 Family and Medical Leave Act (FMLA), which provides 12 weeks of job protected but unpaid leave to some workers.[2] Studies of this policy generally find small or no effects on maternity leave duration, employment, and wages (Klerman and Leibowitz, 1997, 1999; Waldfogel, 1999; Baum, 2003).[3] It is difficult to infer from these studies how job protection would matter as part of a more accessible paid leave policy because the leave is unpaid and the duration of the job protection is very short. Not only do I explicitly show how job protected paid leave affects employment outcomes compared to wage replacement alone, but the policies I study in Great Britain also provide benefits that are much shorter in duration than in most of the countries studied in the existing literature. This is important because, compared to these other countries, Great Britain's policies are much more similar to a policy that could be feasibly implemented in the United States.

This paper is organized as follows: Section 2 provides background on maternity leave in Great Britain. Section 3 provides a simple two period labor supply model that shows how wage replacement benefits and job protection differentially affect employment outcomes. Section 4 describes the data used to evaluate the effects of maternity leave, and section 5 introduces the empirical strategy. The results are discussed in section 6, and some potential mechanisms are shown in section 7. Finally, section 8 provides concluding remarks.

[1]For more details about the FAMILY Act, see `https://www.govtrack.us/congress/bills/114/s786/summary`, accessed September 3, 2016.

[2]Several states have adopted paid family leave policies, which provide approximately six weeks of partial wage replacement to workers. The paid family leave policies in California and New Jersey do not include job protection explicitly, but some workers are simultaneously eligible for FMLA. Rhode Island offer 4 weeks of job-protected paid leave, and New York will start providing job-protected paid leave in 2018.

[3]One reason for these negligible effects is that only about half of all workers are eligible (Ruhm, 1997).

# 2 Background

## 2.1 Literature on the Employment Effects of Maternity Leave Policies

There is a growing literature that looks at the effects of maternity leave policies on employment outcomes. While many studies from a variety of countries clearly show that paid leave has large, positive effects on maternity leave take-up (e.g., Carneiro et al. (2015); Dustmann and Schönberg (2012); Liu and Skans (2010); Rasmussen (2010); Rossin-Slater et al. (2013)), the empirical evidence on employment effects is more mixed. Furthermore, the existing literature cannot separate the effects of having access to wage replacement from the additional effects of job protection. The policy setting in Great Britain allows me to contribute to this literature in two primary ways. First, I use two policy reforms to distinguish between the effects of expanding access to wage replacement benefits and the additional effects of providing job protection benefits. Second, I am able to examine firm-specific employment flows and measures of upward mobility better than in much of the existing work in other countries.

Several cross-country comparative studies find that paid leave is associated with somewhat higher female employment rates (Jaumotte, 2003; Pettit and Hook, 2005). Ruhm (1998) shows that the availability of nationwide leave increases female employment during childbearing years more generally. This could be due to more women working in pre-birth years in order to qualify for paid leave, or due to increased rates of returning to work after the leave ends. The idea that women may work more to qualify for leave benefits if they expect to have a child in the future is consistent with findings from the unemployment insurance literature, which shows that increasing the length of the qualifying period has a significant impact on employment duration (Green and Riddell, 1997).

Studies that exploit policy changes within a country find mixed effects on women's employment outcomes. For example, Baker and Milligan (2008) show that a short maternity leave expansion in Canada increases the proportion of mothers who are employed and on leave, but has no effect on the length of leave taken. However, they do not look at effects on wages or more specific employment outcomes. Gregg et al. (2007) find that maternity leave has large positive effects on the employment rates of married women, but does not significantly affect wages. In

contrast, Ejrnaes and Kunze (2006) find longer amounts of leave have negative effects on both employment and wages for women. Finally, Dahl et al. (2016) study the effects of a series of paid maternity leave expansions in Norway that incrementally increase the period of paid leave from 18 to 35 weeks over six expansions in five years. The authors find no effects of any of the reforms on labor market participation or earnings.[4]

Other studies focus on longer-term employment effects. Lalive and Zweimüller (2009) and Lalive et al. (2013) find that longer amounts of paid leave in Austria increase leave duration, but there are mixed effects on earnings five years after the child's birth.[5] Finally, Schönberg and Ludsteck (2014) find that in Germany, paid leave expansions that exceed the length of job protection have long-lasting negative effects on women's employment outcomes. The empirical literature generally suggests relatively short periods of leave have little effect on wages, but longer leaves may have positive or negative effects. There also may be differential effects across different types of workers. See Hegewisch and Gornick (2011) for a review of the existing literature.

Finally, as mentioned above, maternity leave policies vary by whether or not they include wage replacement and job protection benefits, but also by the duration of these benefits. All three components are likely to affect women's employment outcomes in different ways. While I am able to separate the effects of expanding access to wage replacement benefits from the additional effects of providing job protection benefits, I cannot say anything about the importance of benefit duration. Both Dahl et al. (2016) and Schönberg and Ludsteck (2014) explicitly analyze expansions in maternity leave coverage duration. The evidence suggests that the provision of relatively short periods of paid, job-protected leave may have a stronger impact on long-run employment outcomes than subsequent leave expansions.

[4]Dahl et al. (2016) attempt to identify the effect of wage replacement duration, and the policy variation comes only from changes in the length of the wage replacement above the initial 18 weeks. Eligibility requirements and the amount of job protection during paid or unpaid leave remained fixed. Therefore, the authors do not try to separate the effect of access to leave from the effect of an expansion of benefits. Most mothers in Norway already took leave under the original maternity leave policy as well. Combined with the limited variation in benefits, the margin for employment or other behavioral changes is plausibly smaller than in other settings.

[5]Lalive et al. (2013) use three maternity reforms in Austria to try to identify the relative importance of wage replacement and job protection. However, the reforms only change the length of wage replacement, so they cannot say anything about the effect of job protection in the empirical analysis. Instead, they build a structural search model to simulate the separate effects of job protection and wage replacement. The policy simulations weakly suggest that a policy with both types of benefits improves long-run labor market attachment better than policies with either only job protection or only wage replacement.

## 2.2 Maternity Leave in Great Britain

Great Britain has had official national maternity leave legislation in place since 1976, offering a short period of job protection as well as wage replacement benefits (Waldfogel, 1998). However, until 1993, the period of leave was both short and restricted to those most attached to the labor force. Women were only eligible for coverage if they had worked full time continuously for the same employer for the previous two years or part time for the past five consecutive years. This meant that less than 40 percent of women were eligible. Those who did qualify received a benefit equal to 90 percent of their wage for six weeks and a small flat-rate amount ($£32.85$) for an additional 12 weeks.

In 1994, the work requirements were significantly relaxed so that coverage became almost universal. This 1994 reform (hereafter called the eligibility reform) drastically increased eligibility for wage replacement without changing the amount of benefits offered. Women with 26 weeks of continuous employment (not employer-specific) were eligible for the same wage replacement as before.[6] Women without 26 weeks of continuous employment were eligible for 18 weeks of paid leave at a flat rate, provided they had worked a small set amount in the past 18 months.

In 2000, maternity leave in Great Britain was substantially reformed again. This reform (hereafter called the job protection reform) provided substantial job protection benefits without affecting eligibility or substantially changing the amount of wage replacement available. Starting in 2000, all working women became eligible for one year of job protection during leave, extending the length of job protection beyond the length of paid leave. Paid leave was also extended to six months, but this extension only increased the length of the low flat-rate period. Women who met the minimal work requirements, which essentially amounted to being employed around the time of conception, were reimbursed at 90 percent of their wage for six weeks, but now received a flat-rate amount for up to an additional 20 weeks. Women who did not meet these minimal requirements only received the flat-rate amount for 26 weeks, which was now $£60.20$ per week

[6]Women were eligible for a few weeks of post-birth job protection, however this was largely unenforced (Long, 2012).

(Smith, 2010).[7] In real terms, the flat rate has remained roughly constant over this entire period, and corresponds to only about 30 percent of the mean weekly earnings for women.[8] Because the flat rate benefit is small and the length of the 90 percent wage replacement stayed the same, the main benefit of the 2000 maternity leave reform was the year of job protection. The 1994 reform increased the probability of having received maternity leave wage benefits in the last pay period by 16 percentage points. However, the job protection reform did not further significantly increase the probability of receiving maternity pay. Because eligibility requirements stayed the same after the 2000 reform, this lends support to the idea that the primary benefit was job protection and not the additional weeks of flat rate pay.

Employers are responsible for paying employees during their maternity leave, but 92 percent of the expense is deductible from the employer's National Insurance contribution, which is a mandatory monthly tax/insurance premium paid to the government.[9] Because employers can deduct most of the wage replacement from their insurance payment, the direct costs of the policy to the the employer are minimal.[10] National Insurance provides pension, unemployment, illness/disability, maternity, and bereavement benefits to all workers, and therefore information about benefit structures and eligibility requirements is relatively easy to disseminate.

The structure of Great Britain's maternity leave policy reforms allow me to study both the extensive and intensive margins of nationally provided maternity leave benefits. The 1994 eligibility reform drastically increased eligibility for wage replacement without changing the amount of benefits offered, and the 2000 job protection reform provided substantial job protection benefits without affecting eligibility or substantially changing the amount of wage replacement available. This allows me to identify the relative importance of wage replacement benefits and job

[7] In 2007, the entitlement was increased to 39 weeks, with 6 weeks at 90 percent of their wage and the remaining 33 weeks at a flat rate. This flat rate is $£139.58$ in 2016.

[8] Average weekly earnings data comes from the Office for National Statistics: `http://www.ons.gov.uk/employmentandlabourmarket/peopleinwork/earningsandworkinghours/datasets/allemployeesashetable1`.

[9] Most workers are covered by Statutory Maternity Pay, which is directly paid by the employer but then is deductible from the employer's National Insurance contributions. All workers 16 and over who meet a minimum earnings requirement make National Insurance contributions as well. Self-employed individuals and others unattached to a particular employer (but who qualify for leave benefits) are paid directly by the Department of Social Security. Employees can start their maternity leave as soon as 11 weeks before the child's expected due date. Employers must be notified at least 21 days before the intended start date. The 90 percent wage replacement is based on 90 percent of one's average weekly pre-tax earnings in the 12 weeks before the start of maternity leave.

[10] Other costs to employers are discussed in section 7.

protection benefits on employment, wages, and other measures of career success more precisely than has been done in the existing literature.

# 3 Theoretical Framework

Although increasing maternity leave benefits should increase leave-taking in the short run, the effect of maternity leave policies on long-term employment outcomes is theoretically ambiguous (Klerman and Leibowitz, 1994; Han et al., 2009). In this section, I provide a two period labor supply model to show how maternity leave policies can affect employment decisions in the period that the benefit is received as well as in future periods. I start by analyzing labor supply decisions in this model when no maternity leave is available, and then consider the effects of two different policies: one that provides wage replacement benefits only, and another that provides job protection in addition to wage replacement. Evaluating how labor supply decisions change under these three scenarios yields predictions about the effects of these maternity leave policies that I can test empirically.

## 3.1 Two Period Labor Supply Model

Consider a simple model where women gain utility from consumption ($c_t$) and childrearing ($k_t$). They are endowed with a stock of time $T_t = 1$ in each period that can be allocated between market work ($h_t$) and staying home to care for children, and they choose between work and childrearing to maximize the present value of lifetime utility over two periods ($t = 0, 1$), subject to a budget constraint.[11] The utility function is linear in consumption and childrearing such that in each period a woman chooses to spend all her time either working or childrearing ($k_t = 0$ or $k_t = 1$, respectively).[12] Women differ in how much they value childrearing compared to consumption, and $\frac{\sigma}{1-\sigma}$ represents the individual's marginal rate of substitution between the two. The parameter $\sigma \in [0, 1]$ is constant for a given woman over time, and is drawn from a

[11] Individual subscripts are omitted for notational simplicity.

[12] One can also interpret this assumption to mean that employers offer rigid wage contracts that require full-time work in a given period. This assumption is standard in models of the effect of paid sick leave policies on absences in both static (Pichler and Ziebarth, 2016; Stearns and White, 2016) and dynamic (Brown, 1994; Barmby et al., 1994) settings.

probability distribution.[13]

Women face a potential wage $w_0$ in period 0, which they receive if they choose to engage in market work. The wage a woman can earn in period 1 is a function of her labor supply decision in period 0. If she works in period 0 ($k_0 = 0$), then her potential wage in the next period is the same.[14] However, she faces a wage penalty, $\delta w_0$ with $0 < \delta < 1$, for periods of non-work such that $w_1 = w_0(1 - \delta k_0)$. This penalty can be interpreted as a loss in firm-specific human capital, the search cost of finding a new job, or other barriers to entry into the labor market. Individuals have access to credit and can borrow and lend at rate $r > 0$, and have a rate of time preference, $\rho$.

Finally, fertility is exogenous and women are assumed to experience a birth at the start of the initial period. While fertility is not exogenous in my empirical setting, I show in Section 5 that neither of the policy reforms I study affect fertility. Women who have access to maternity leave receive a fraction of their potential wage, $b_0 w_0$, if they stay home and engage in childrearing rather than participate in market work. For simplicity, eligibility for maternity leave does not depend on previous work experience. I assume the maternity leave wage replacement rate, $b_0$, is strictly less than one ($0 < b_0 < 1$). This captures the possibility that there are non-monetary costs to not working, even if there is full wage replacement. Women receive non-wage income $I_t$ in each period regardless of their labor supply decision, but do not receive maternity leave benefits in the second period.[15]

[13]The model can be generalized to allow a woman to have a different value of $\sigma$ in each period. This would allow for mothers to value time spent with infants differently from time spent with toddlers, for example. If both values are known at the beginning of period 0 when the woman makes the decision whether or not to work, then the model is solved as described below and the indifference points do not change. However, if $\sigma_1$ is unknown at the beginning of period 0, then the mother will be indifferent to working or childrearing in period 0 when the utility of working in period 0 plus the expected utility in period 1 conditional on working in period 0 is equal to the utility of childrearing in period 0 plus the expected utility in period 1 conditional on childrearing in period 0. I extend the model to include this possibility in Appendix A, and the main predictions remain the same.

[14]The main predictions of the model do not change if I allow for wage growth over time.

[15]One can think of $I_t$ as spousal earnings or public benefits, and $I_t > 0$ ensures that consumption is strictly positive. For simplicity, the level of $I_t$ is assumed to be independent of the labor supply decision. However, the results hold if this assumption is relaxed to allow $I_t$ to be increasing in $k_t$. This could be the case if parents jointly make labor supply and childcare decisions or if public benefit receipt is decreasing in wage income.

Women maximize the present value of their lifetime utility subject to their budget constraint:

$$\max_{k_t} \sum_{t=0}^{1} \beta^t [(1-\sigma)c_t + \sigma k_t] \tag{1}$$

$$\text{such that } c_0 + \gamma c_1 = I_0 + (1-k_0)w_0 + b_0 w_0 k_0 + \gamma I_1 + \gamma(1-k_1)(w_0(1-\delta k_0))$$

$$h_t + k_t = 1$$

where $\beta^t = \frac{1}{(1+\rho)^t}$, $\gamma = \frac{1}{(1+r)^t}$, and the other variables are as defined above. For simplicity, I assume $\beta = \gamma$.[16]

Mothers decide to work ($W$) or stay home ($K$) in each period in order to maximize the present value of their lifetime utility, where the present value of utility resulting from each pair of possible decisions is given as follows:

$$U(W_0W_1) = (1-\sigma)[I_0 + \gamma I_1 + w_0 + \gamma w_0]$$

$$U(K_0W_1) = (1-\sigma)[I_0 + \gamma I_1 + b_0 w_0 + \gamma w_0 - \gamma\delta w_0] + \sigma$$

$$U(W_0K_1) = (1-\sigma)[I_0 + \gamma I_1 + w_0] + \sigma\gamma$$

$$U(K_0K_1) = (1-\sigma)[I_0 + \gamma I_1 + b_0 w_0] + \sigma + \sigma\gamma$$

It is then possible to solve for the threshold values of $\sigma$, $\sigma_t^*$, where a woman is indifferent between working and childrearing in each period. If $\sigma > \sigma_t^*$, the woman will choose to stay home in that period; otherwise she will choose to work. Assuming a distribution of $\sigma$ across women, analyzing how these indifference points shift in response to changes in the other model parameters directly corresponds to predictions about how the share of women working versus staying home to care for children will change.

I next consider the work and childrearing choices a woman will make under three different policy scenarios. First, I explain what will happen when a mother does not have access to a maternity leave policy. I next introduce a policy that provides wage replacement benefits in the period the child is born. Finally, I consider a maternity leave policy that provides job protection

[16]This assumption, or the assumption that individuals cannot borrow or save across periods, is standard in this literature. See, for example, Brown (1994); Barmby et al. (1994).

benefits in addition to wage replacement, where job protection eliminates the wage penalty for not working. These three scenarios mirror the policy changes that I study empirically.[17]

### 3.2 No Maternity Leave

First, consider the case when no maternity leave benefits are available so that $b_0 = 0$. When $b_0 = 0$, a woman will choose to either work in both periods or stay home in both periods.[18] The indifference point between work and childrearing is the same in both periods: $\sigma^*_{N0} = \sigma^*_{N1} = \frac{w_0}{1+w_0}$, where the subscript $N$ denotes the no maternity leave case. If $\sigma < \frac{w_0}{1+w_0}$ then the mother will work in both periods; otherwise she will stay home in both periods.[19] This indifference point is increasing in $w_0$, so assuming a distribution of $\sigma$ across mothers, as wages increase more women will participate in market work and fewer will choose to stay home. The first line of Figure 1 shows the optimal work decisions in each period for women with different values of $\sigma$.

### 3.3 Maternity Leave with Wage Replacement

Now suppose that a maternity leave policy provides wage replacement benefits in period 0 if the mother does not work such that $0 < b_0 < 1$. Because this lowers the opportunity cost of childrearing in period 0, women are more likely to do so. More formally, a woman will always choose to stay home in period 0 if her value of $\sigma$ is greater than $\sigma^*_{R0} = \frac{w_0(1-b_0+\gamma\delta)}{1+w_0(1-b_0+\gamma\delta)}$, where the subscript $R$ denotes the case of maternity leave with wage replacement only.[20] Then $\frac{\partial \sigma^*_{R0}}{\partial b_0} < 0$. As the maternity leave wage replacement benefit rate increases, the threshold at which mothers choose to stay home moves to the left. In other words, all else equal, more mothers choose to stay home (and fewer choose work) in period 0 as wage replacement benefits increase. The

[17] In the empirical analysis, I study two policy reforms that modify an existing maternity leave policy. The first is a policy that expands eligibility, holding the existing level of benefits constant. This is akin to going from no maternity leave policy to one with wage replacement benefits for the women who gain access, and no policy change for the subset of women who already have access. The second reform provides job protection to all eligible women.

[18] Because of the wage penalty $\delta$, conditional on childrearing in period 0, the utility from working in period 1 is strictly lower than the utility from working in period 0. Therefore, if a woman chooses to stay home in period 0, she will always choose to stay home in period 1 as well. Similarly, conditional on working in period 0, the utility from the labor supply decision in period 1 is the same as in period 0. Thus, if a woman chooses to work in period 0, she will also do so in period 1.

[19] See proof in Appendix A.

[20] This is a necessary condition in the likely case that $b_0 > \delta(1+\gamma)$. If $b_0 < \delta(1+\gamma)$, then it is a sufficient, but not necessary condition for choosing not to work in period 0. See Appendix A for more details.

indifference condition also shows that the threshold is increasing in the wage penalty rate $\delta$ ($\frac{\partial \sigma^*_{R0}}{\partial \delta} > 0$). This means as the future discounted opportunity cost of not working rises, women are less likely to take leave in period 0. Additionally, $\frac{\partial^2 \sigma^*_{R0}}{\partial \delta \partial b_0} > 0$. This means that increasing wage replacement benefits will be less effective at increasing childrearing rates in period 0 when the professional cost of work absences is high. Although not modeled here, it is reasonable to think that this cost is higher for higher skilled women. As suggested by Anderson et al. (2002), a large part of this cost is likely driven by the loss of firm-specific human capital. Lower-skill workers are likely to be less affected by this loss.

Because wage replacement benefits are only received in period 0, they do not directly influence employment decisions in the subsequent period. However, increasing $b_0$ increases the probability of not working in period 0, and the decision of whether or not to work in period 0 affects the potential wage faced in period 1. Therefore, the threshold for not working in the second period, $\sigma^*_{R1} = \frac{w_0(1-\delta k_0)}{1+w_0(1-\delta k_0)}$, is decreasing in $b_0$ because $\frac{\partial \sigma^*_{R1}}{\partial k_0} < 0$. Because of the wage penalty for not working in period 0, increasing the wage replacement rate decreases the share of mothers working in the later period even though there is no additional maternity leave benefit in the later period. Thus, increasing $b_0$ decreases the share of women working in both periods. As long as the wage replacement rate is greater than the discounted wage penalty rate ($b_0 > \delta(1+\gamma)$), the threshold for childrearing in period 0 will be to the left of the threshold for childrearing in period 1 and some women will choose to only work in period 1.[21] This is shown in the second line of Figure 1.

### 3.4 Maternity Leave with Wage Replacement and Job Protection

Finally, suppose that a maternity leave policy includes job protection in period 0 in addition to the wage replacement benefits. Job protection is modeled such that the wage penalty for not working in period 0 is eliminated ($\delta = 0$).[22] This means that the labor supply decision in one

[21] When $b_0 < \delta(1+\gamma)$, women will always choose to either work in both periods or not work in both periods. See Appendix A for more details.

[22] This assumption can be relaxed such that job protection decreases $\delta$ but $\delta > 0$. Then the indifferent points for working in periods 0 and 1 are the same as in the case of maternity leave with wage replacement only. As explained above, decreasing $\delta$ will induce more women to not work in period 0. Decreasing $\delta$ will also cause more women to work in period 1 because $\frac{\partial \sigma^*_{R1}}{\partial \delta} \leq 0$. However, these effects will be less extreme than the case where job protection

period is now independent of the labor supply decision in the other period. Women will be indifferent between work and childrearing in each period when:

$$\sigma^*_{J0} = \frac{w_0(1-b_0)}{1+w_0(1-b_0)}$$
$$\sigma^*_{J1} = \frac{w_0}{1+w_0}$$

where $J$ indicates the maternity leave policy with job protection as well as wage replacement benefits. If $\sigma > \sigma^*_{Jt}$, the woman will choose to stay home in that period; otherwise she will choose to work.[23] This is shown in the third line of Figure 1.

Similar to the case of wage replacement only, $\frac{\partial \sigma^*_{J0}}{\partial b_0} < 0$. This means that as the maternity leave wage replacement rate increases, the threshold at which mothers choose to stay home moves to the left. In other words, all else equal, more mothers choose to stay home. Furthermore, $\sigma^*_{J0} < \sigma^*_{R0}$. This means that a maternity leave policy with job protection will increase the childrearing rate in period 0 more than a policy with wage replacement only. However, with job protection, employment in the future period will increase relative to a policy with wage replacement only because $\sigma^*_{J1} > \sigma^*_{R1}$.

### 3.5 Policy Implications

By comparing the optimal work and childrearing decisions across the three policy environments presented above, this model yields several testable predictions about the effects of maternity leave policies on short and long term employment decisions:

1. The introduction of a maternity leave policy that provides wage replacement benefits but no job protection will decrease the share of mothers at work in the period the child is born (and increase the share of mothers at home).

2. The introduction of a maternity leave policy that provides wage replacement benefits but no job protection will also decrease the share of mothers at work in future periods.

equates to $\delta = 0$.

[23] More specifically, the mother will work in both periods when $\sigma < \sigma^{*0}_J$, will stay home in the period 0 and work in period 1 when $\sigma^*_{J0} < \sigma < \sigma^*_{J1}$, and will stay home in both periods when $\sigma > \sigma^*_{J1}$. See Appendix A for more details.

3. Holding the wage replacement rate constant, introducing job protection will further decrease the share of mothers at work in the period the child is born.

4. Job protection will increase the share of mothers at work in future periods, compared to an otherwise identical policy without job protection.

These predictions are summarized in Figure 1.[24] The theoretical results from this section illustrate how these two different dimensions of maternity leave benefits can affect immediate and longer-term employment outcomes. The analysis that follows empirically tests these predictions by examining the effects of a maternity leave expansion that substantially increases access to maternity leave wage replacement benefits, and a second expansion that provides women with job protection benefits without changing eligibility requirements or the wage replacement rate.

# 4 Data

The data used in this analysis comes from the British Household Panel Survey (BHPS). This longitudinal survey interviews individuals every year from 1991 to 2009.[25] The BHPS consists of a nationally representative sample of individuals and households recruited in 1991, and each adult in the household is surveyed individually. Over the 18 survey waves, there are over 338,000 observations from 32,342 individuals. The survey asks a wide variety of questions about household composition, education and training, labor market activity, and income. In particular, this survey focuses on labor market spells, tracking the length of time in one's current job (or non-employment spell), firm characteristics, wages, and other income sources from government programs such as maternity leave payments. The survey also contains the month and year of birth of all the respondent's children, and information about whether or not the mother is currently on maternity leave. Particularly relevant to my analysis, the data allows for the construction of complete employment histories as well as fertility histories. Unlike in other studies, I am able to analyze job-specific employment flows, not just employment status. For example I

[24] Assuming a distribution of values of $\sigma$ across mothers, the share of mothers working in period 0 is given by the blue area, and the share of mothers working in period 1 is the sum of the blue and red areas.

[25] The new UK longitudinal household study, Understanding Society, was started in 2009. The majority of BHPS respondents were integrated into this yearly survey starting in 2010 so it is possible to continue to follow them past 2009. However, I am not using this data due to concerns about selective attrition.

can see if a woman continues in the same job after maternity leave, or if she starts a new job or exits the labor force.

The survey also allows me to construct a proxy for maternity leave eligibility using each individual's employment history. Although the survey starts in 1991, very specific employment and fertility histories can be traced back further for most individuals. Importantly, one can recover most of the key variables used in this analysis for any month-year in the respondent's adult life.[26]

The majority of this analysis limits the sample to women with young children, and compares women who have infants to women with slightly older children age 3-4. However, other control groups including mothers of younger children, mothers of older children, women without children, and fathers of infants are used to check the robustness of the results.

Table 1 shows summary statistics for the whole BHPS sample as well as for the sample of mothers, mothers of infants, and mothers of youngest children age 3-4. Mothers in Great Britain are much less likely to be married than women in the United States, and are more likely to be cohabiting. The sample is predominantly white, and about eight percent of mothers have an infant. About 70 percent of mothers are employed and about six percent are absent from work at any given time.

## 5 Empirical Strategy

To identify the effect of maternity leave policies on employment outcomes, I employ a difference-in-differences (DD) framework with individual fixed effects. This model compares changes in outcomes of mothers who give birth before and after the maternity leave reforms to corresponding differences in the comparison group of women with slightly older children, who are unlikely

[26] It is therefore possible to extend the panel backwards to years prior to 1991. As is explained below, my preferred specification limits the sample to short windows of time around each policy change, so doing so is not necessary. None of the main results use reconstructed observations, but the pre-1991 trends in employment and fertility rates that are constructed from these histories are comparable to the trends found in the cross-sectional General Household Survey.

to be directly affected by the policy changes.

$$Y_{i,t+b} = \alpha + \beta_1 Infant_{it} + \beta_2 Eligibility_{it} + \beta_3 JP_{it} + \beta_4 Eligibility * Infant_{it} + \beta_5 JP * Infant_{it} + X'_{i,t+b}\gamma + \delta_i + \theta_t + \varepsilon_{i,t+b} \quad (2)$$

The dependent variable, $Y$, is the employment outcome of individual $i$, measured $b$ years after the reference year $t$. The binary variable $Infant$ is equal to one if individual $i$ has an infant less than one year old in year $t$ and zero otherwise. The $Eligibility$ indicator takes a value of one in years after the 1994 maternity leave reform takes effect. The 1994 eligibility reform expanded the population of women who were eligible to receive maternity leave benefits but did not change the value of the available benefits. The 2000 job protection reform substantially increased the amount of job protection available but did not change the eligibility requirements. Therefore, $JP$ is an indicator for years 2000 and later when this policy is in effect. The vector $X$ contains individual level time-varying controls including dummies for age, marital status, educational attainment, number of kids, and the age of the youngest child in the household. Finally, $\delta_i$ and $\theta_t$ are individual and year fixed effects, respectively. Standard errors are clustered at the individual level.

The coefficients of interest are $\beta_4$ and $\beta_5$. These are the DD coefficients representing the effect of the two maternity leave reforms on employment outcomes. The first, $\beta_4$, is interpreted as the effect of the eligibility reform on maternal employment outcomes $b$ years after the birth of a child. The second, $\beta_5$, is the additional effect of providing job protection benefits, holding constant eligibility requirements and wage replacement benefits.

Several potential issues arise when estimating the DD model described above. The key identification assumption is that changes in the trends of employment outcomes would have been the same for the treatment and comparison groups in the absence of the maternity leave policy changes. Although this assumption is fundamentally untestable, I mitigate concerns about differences in trends in several ways. First, the results are robust to the use of multiple alternative comparison groups. For reasons described below, the main comparison group is mothers with youngest children aged 3-4 in year $t$, but the results are robust to changes in the minimum and

maximum child age thresholds for inclusion in the control group. Results are also qualitatively similar if women without children or fathers of infants are used as the control group.

Second, in most specifications the treatment effects are identified with only three years of data on either side of each policy change. Restricting the analysis to include fewer years reduces concerns about differential trends in the treatment and control groups. This restriction is also necessary to look at long-term employment outcomes. The DD coefficients in Equation (2) identify the effect of each policy change under the assumption that the control group is not affected by the policy. If mothers in the control group were exposed to the relevant policy in previous years, the DD coefficients will be biased toward zero.[27] By limiting the analysis to a three year window around each policy change, I can ensure that no one in the control group is ever exposed to the relevant treatment as long as they do not have children younger than age three.[28] Therefore, the main control group used to look at employment outcomes consists of mothers with youngest children 3-4 years old.

Finally, Table 2 shows that the policies are uncorrelated with differential changes in fertility or demographic characteristics between mothers of infants and mothers of 3-4 year olds. This is an indirect test of the key identification assumption, and each column shows a separate regression that estimates the effect of the maternity leave policies on observable maternal characteristics using the main empirical specification (Equation 2). The first column shows that neither maternity leave policy reform affects fertility.[29] The remaining columns show there are no differential trends in maternal age, marital status, education, or race that are correlated with the policies. The single exception is a significantly lower share of Asian mothers of infants after the job protection reform. However, there is not a significant difference in the maternity leave rate among Asian and white mothers.

[27]This will cause attenuation bias under the assumption that the sign of the effects on the treatment and control group are the same, as is likely the case here. This issue is not a concern when looking at short-term outcomes. For example, a maternity leave policy should not affect the probability that a mother whose youngest child is three is currently on maternity leave, even if the policy did cause her to take maternity leave when the child was born.

[28]This strategy is feasible because the effects of the policy changes were fairly immediate. Because the national maternity leave policy has existed in Great Britain since 1976, the benefits are common knowledge and information about the policy changes was easy to disseminate. This strategy would not work if take-up of the policy was low in the first few years. See Figure 2 for evidence that suggests increased take-up of both policy reforms was fairly immediate.

[29]This regression is estimated using a simple difference model. All other columns estimate the DD model.

# 6 Results

## 6.1 Effects on Leave-Taking

Figure 2 shows an event study of the probability of being on leave at the time of the survey for all years of data.[30] This figure plots the difference between the leave-taking rate among mothers of infants compared to mothers whose youngest child is 3-4 years old in each year, with the coefficients normalized to zero in 1994. There is no evidence of an upward trend in leave-taking prior to the introduction of the eligibility reform in 1994. Take-up of both reforms is immediate and relatively stable in subsequent years. Again, there is no evidence of an upward trend between the 1994 and 2000 reforms. This is important because in the subsequent regression analysis, treatment effects are identified with only three years of data on either side of each policy change. There are two reasons why the policy reforms appear to increase leave-taking so immediately. First, information about the reforms was relatively easy to disseminate because maternity leave falls under the National Insurance system which covers most paid leave and other benefits for workers. Second, the reforms both went into effect near the beginning of the calendar year, and the BHPS surveys respondents in the last quarter of the year. Most respondents therefore had over 9 months to learn about the reform between its implementation and their survey date.

Table 3 shows the DD effects of both the maternity leave eligibility and job protection reforms on the probability of being on leave and the probability of being employed in the current period ($b = 0$ in Equation (2)). Column (1) suggests that both reforms cause a substantial increase in leave-taking of about 10 percentage points. The increase in maternity leave take-up from the eligibility reform is driven by an increase in the number of mothers who take leave, whereas the job protection expansion increases the length of time that mothers are on leave. Because individuals are only surveyed once a year in the BHPS, increasing the length of leave a mother takes increases the probability that she is absent from work at the time of the survey. As the model in Section 3 predicts, providing access to maternity leave wage replacement benefits increases the share of mothers of infants who are on leave, and adding job protection benefits further in-

[30] The event study looks similar when using other control groups as well. Because the probability of being on leave is very low for mothers of older children, it also looks very similar when plotting the raw leave rate of mothers of infants.

creases the leave-taking rate. In fact, both policy reforms roughly double the share of mothers of infants who are on leave. The magnitude of these effects are similar to those found in the U.S. and Canada (e.g., Rossin-Slater et al. (2013); Baker and Milligan (2008)).

Column (2) of Table 3 shows the effects of the maternity leave policies on the probability of being employed at the time of the survey. Unlike leave-taking, the eligibility reform has no significant effect on the probability that a mother of an infant is employed at the time of the survey, although the point estimate is negative. However, the job protection reform has a substantial positive effect on the probability that a new mother is currently employed. Providing job protection increases the probability of being employed by 3.8 percentage points. This represents about a 7 percent increase in the probability that a new mother is employed. Because job protection requires that employers allow workers to go on extended periods of leave, it makes sense that it has a large positive effect on employment in the short-run.[31] Job protection allows mothers to remain employed when they go on leave even if they do not expect to return to work. When women can keep their job, the transition back into work is also less costly, which may induce some women to return to work when they may otherwise not have.

These results are robust to a number of alternate specifications and control groups. Appendix Tables B1 and B2 show that the effects of the two reforms on the probability of being on leave and being employed at the time of the survey are not sensitive to excluding individual level controls or maternal fixed effects or changing the choice of control group.[32] Using fathers of infants as a control is not my preferred specification because there may be spillover effects of maternity leave within the household, which would bias the results. Using mothers of youngest children age 1-2 as a control group is fine when looking at short-term outcomes as in these two tables, but will attenuate effects when looking at longer-term outcomes because some of the control group mothers would also be treated. Finally, it seems more likely that the common trends

[31]Consistent with this result, the job protection reform has a small but statistically significant positive effect on self-reported job satisfaction in the year the child is born. This effect is entirely driven by an increase in job satisfaction with respect to job security. However, there are no effects of either policy on overall job satisfaction or satisfaction with job security in future years. These results are available upon request.

[32]The results are stable when using mothers with a youngest child age 1-2, mothers with a youngest child under age 15, women without children, and fathers of infants as the control group. Estimates of difference-in-difference-in-difference effects, comparing mothers of infants to mothers of youngest children age 3-4, relative to the corresponding difference between fathers of the same age children, before and after the introduction of each maternity leave reform, are similar as well.

assumption is satisfied when comparing mothers of infants to mothers of young children than when comparing them to women without children, mothers of much older children, or fathers. I use mothers of youngest children age 3-4 as my preferred control group for this reason.[33]

The increased availability of job-protected leave should increase short-run employment rates somewhat mechanically. Even if women do not change their behavior in the long run as a result of the policy, job protected leave allows them to delay quitting their jobs until the end of the maternity leave period. Some of this increase likely comes from women who delay quitting their job. Part of the increase may also be driven by an increase in the number of women who return to work soon after their maternity leave ends. For women who plan on returning to work, the ability to return to one's same employer reduces search costs and matching frictions. Therefore, it is important to look at longer-term effects on employment to determine the extent to which maternity leave policies are effective at improving women's labor force attachment.

### 6.2 Effects on Long-Run Employment

Table 4 shows the effects of each maternity leave reform on the employment outcomes of new mothers 1, 3, and 5 years after giving birth. The comparison group consists of mothers with youngest children 3-4 years old in year $t$. The sample is limited to three years on either side of each reform, so mothers in the control group were never exposed to the relevant policy. Columns (1)-(3) show that the two maternity leave reforms have different effects on medium and long-term employment rates.[34] The eligibility reform has substantial effects on the probability of being employed one to three years later, but no lasting effects on employment after that. The policy increases employment by 6.5 percentage points (10.1 percent) one year later and 3.7 percentage points (5.8 percent) three years later. On the other hand, providing substantial job protection benefits has longer-lasting effects on employment. The job protection reform increases the probability of being employed in three years by 3.9 percentage points (7 percent) and the probability of being employed in five years by 4.5 percentage points (8 percent). Job

[33] Because the treatment effects are identified off of just three years of data on either side of each policy change, however, differential trends may be less of an issue than in other DD settings.

[34] Appendix Table B3 shows that the effects of the maternity leave reforms are also robust to using different control groups. The effects are qualitatively similar when using women without children, mothers with a youngest child age 3-15, and fathers of infants as control groups.

protection does not significantly affect employment one year after giving birth when eligibility is held constant. Columns (4)-(6) limit the sample to married mothers. This is because a substantial fraction of the sample is unmarried, and single mothers are more likely to be in the labor force than are married mothers. If single mothers have to work to support their families, their long-run employment outcomes should be less affected by maternity leave policies. The DD coefficients are only very slightly larger in magnitude, but because married mothers are less likely to be working, the effects are larger in percentage terms. These effects are reasonably consistent with other estimates in the existing literature. Ruhm (1998) uses data across nine European countries to estimate that if women were given access to 40 weeks of job protected paid leave, employment rates would rise by 7-9 percent. Lalive and Zweimüller (2009) find that doubling the length of job protected paid leave from 12 to 24 months in Austria increases the probability of being employed 3-10 years after a birth by 5-9 percentage points, although their estimates are noisy.

These results imply that providing job protection is important in increasing long-term employment rates of mothers. However, one concern is that while the eligibility reform only directly affected mothers who were not already eligible for maternity leave wage replacement benefits, all working mothers gained access to job protection under the job protection reform. To rule out the possibility that the greater effects of the job protection reform on long-term employment are driven by differences in the population of mothers affected by each reform, I create a measure of pre-reform eligibility for each mother based on her work history as of 1993. Mothers who would have been eligible for maternity leave wage replacement benefits under the original maternity leave policy are classified as "Already Eligible" and mothers who did not meet that criteria in 1993 are assumed to have potentially become eligible under the 1994 eligibility reform ("Became Eligible"). Table 5 shows the differential effects of the two policy reforms on these groups. While only the "Became Eligible" group should have been affected by the eligibility reform, both groups were given access to job protection under the job protection reform. The table shows, as expected, that the effects of the eligibility reform on leave taking and employment 0-5 years later are driven by those in the "Became Eligible" group, who should have been directly affected by the policy change. While noisy, the additional effects for the "Already Eligible" group

are all of the opposite sign and similar in magnitude. The total effects on employment outcomes for this group are close to zero and insignificant.

On the other hand, the results for the job protection reform show that the effects are the same for both groups. The effects for the individuals who likely became eligible for wage replacement under the 1994 reform are generally consistent with the overall results, and the additional effects for the group who were already eligible for wage replacement benefits in 1993 are all very close to zero, suggesting that the total effects are the same for both groups. This means that the differential effects of the two policy reforms are driven by the different dimensions of benefits offered and not by differences in the population of mothers who gained access to those benefits.

### 6.3 Effects on Hours Worked and Earnings

If maternity leave causes women to try to balance both work and childcare, then even if employment rates increase on the extensive margin, average hours worked may fall. Table 6 shows the effects of each maternity leave reform on average usual hours worked per week. The first four columns show that, conditional on working, women are choosing to work less after job protection becomes available. Mothers work 2.6 hours less per week on average in the year the child is born, which corresponds to a 10.8 percent decrease from a mean of 24 hours per week. In the year after the child is born, women still work 12.9 percent fewer usual hours per week, but the effect on hours worked fades over time. Conditional on employment, women work the same amount on average as they did before the job protection reform by three years after the birth of a child. Columns (5)-(8) show the effects on an unconditional measure of hours worked. The higher employment rates associated with the job protection reform largely negate the negative effect seen in the first four columns. Finally, Table 7 shows that the mothers who now work when job protection is available are less likely to work full-time. Consistent with Columns (1) and (2) of Table 6, this effect is only significant up to one year after the child is born, and then fades over time.

These maternity leave reforms may have effects on income as well. If the policies ease the transition back into work after taking time off to care for a new infant, then access to maternity leave policies could be expected to increase wage income. On the other hand, if the policies

cause lower ability women to enter the workforce or reduce their usual hours of work (as shown in Table 6), average earnings could fall. Figure 3 shows how the monthly income distribution of mothers of infants changes as a result of the two maternity leave reforms. The figure shows the DD coefficients for each policy reform on the probability of being in each decile of the usual monthly earned income distribution, and the error bars give the 90 percent confidence interval for each point estimate. The coefficients come from ten separate regressions using the same specification as described above, again using mothers of children age 3-4 as the control group.[35] The income distribution includes all women regardless of employment status, so the bottom two deciles include women with zero earned income and who are not working.

Figure 3 shows that, for the most part, the eligibility reform does not have any significant effects on earnings in the year the child is born. The only significant DD coefficient is on the probability of being in the third decile, which indicates that some mothers are shifting from not being employed to working part time. The average woman in the third decile works only 15 hours per week and earns less than $£350$ per month. This is consistent with the idea that women may be taking part-time, perhaps short-term or low-skill, jobs in order to qualify for the maternity leave benefits.[36] However, the small and imprecise coefficients at all other points in the distribution suggest that any short term effects of the eligibility reform work through the extensive margin. On the other hand, consistent with Table 3, the figure shows that the job protection reform increases employment rates among new mothers. This can be seen by the negative and significant effects on being in the bottom quintile of the income distribution. Interestingly, this reform has significant effects on income throughout the distribution. The shift into employment appears to be distributed over the middle 50 percent of the income distribution, while there is some evidence of a relative decline in the share of top earners. This is consistent with the idea that the policy causes some high ability women to remain in the labor force when

[35]The income deciles are calculated using all mothers age 21-45, but the regressions are limited to mothers of young children. The results are qualitatively similar but noisier if deciles are calculated using only the women in the sample. Income is converted to year 1995 £. The top 1 percent of earners are excluded.

[36]The variable used in this analysis measures usual monthly income, which can be averaged over months employed and not employed. Some of the women in this decile are not currently employed at the time of the survey and have an infant, suggesting the DD coefficient could be measuring changes in employment prior to giving birth, which would be consistent with the women trying to qualify for leave benefits.

they otherwise would not have, but to reduce the amount of hours that they work.[37]

Figure 4 shows the analogous DD effects of the two policy reforms on mothers' earnings five years later. Again consistent with the results in Table 4, there is no effect of the eligibility reform on the probability of being employed (i.e., in the first decile of the income distribution), but the job protection reform has a positive effect on long-run employment. What is most interesting is that there are large, significant negative effects of the job protection reform on the probability of being at the top of the income distribution. As will be shown in the next section, these effects are consistent with decreases in the probability of entering into management positions.

Finally, Table 8 shows the effects of the eligibility and job protection reforms on mean usual monthly work income in five years. The first column shows the effect when both employed and not employed women are included in the sample (as in Figure 4). As can be inferred from the figure, neither policy has a significant effect on long-run mean earnings. Column (2) conditions on being employed five years after the reference year, so all women in the sample have positive earnings. With this restriction, the job protection reform has a negative and statistically significant (at the 10 percent level) effect on long-run wage of about $£71$ per month, from a mean of $£1285$. This represents a five percent decline in average monthly earnings from the sample mean, and appears to be driven by a reduction in the earnings of top-earners.

The results in this section indicate that both the wage replacement and job protection reforms increase maternity leave take-up. Yet only the latter reform has a lasting effect on employment, which suggests that job protection benefits are a crucial part of any maternity leave policy intended to increase female labor force participation rates. Even though the job protection reform increases employment on the extensive margin, it causes women to work fewer hours in the first two years after the child is born. Combined with the increase in leave-taking at birth, this could explain why, despite the increase in the employment rate, I find no positive effects on long-term earnings. Although there is little effect on mean monthly earnings measured five years later

[37] There is little effect on the distribution of hourly wages, conditional on being employed at the time of the survey. While there is a significant increase in the share of new mothers in the lowest decile of the distribution for the job protection expansion, all other effects are small and very noisy. This is partially due to the fact that the hours worked variable may not be very reliable, and contains a significant amount of bunching at 20, 30, 35, and 40 hours per week. Another issue with this measure is that if women temporarily change their hours worked around the time of birth but are salaried, hourly wage would be very noisy.

overall, there is about a five percent decline conditional on working, and these effects are concentrated at the top of the earnings distribution.

# 7 Mechanisms

## 7.1 Heterogeneity

The results presented so far suggest that although job protection can increase the long-term employment rate of mothers, it does not appear to close the gender gap in earnings or hours worked. When thinking about the policy implications of these results, it is important to understand whether they are driven by particular groups of women. This section explores heterogeneity in the effects of wage replacement and job protection benefits by the mother's subsequent fertility and by the mother's educational attainment.

One of the most important predictors of whether or not a mother is working is the age of her youngest child. Table 9 replicates the first three columns of Table 4, but splits the sample by whether or not the mother has another child in the future.[38] This is possible because neither policy affects the probability of having a subsequent birth. Mothers who will go on to have more children may make different employment decisions than mothers who do not. The results for the maternity leave eligibility reform suggest expectations about future childbearing are important.[39] The eligibility reform has significant one to three year effects on the probability of being employed, but only for mothers who eventually have another child. This suggests that mothers who expect to have another child go back to work in order to re-qualify for the maternity leave benefit. The majority of siblings are born less than four years apart, so the fact that there are only significant effects for the first three years after birth is consistent with this idea.

The effects of the maternity leave job protection expansion are larger for women who do

[38] Because the BHPS ends in 2009, I am less likely to have complete fertility information for mothers who give birth at the end of my sample period than for women who give birth in the early 1990s. To avoid potential bias related to this issue, I only look at the six years after the birth. If a woman has another child within six years, she is categorized as having a subsequent child. If a woman does not have another child within six years, she is categorized as not having more children. The results are not sensitive to making this restriction or to using a five year cutoff.

[39] The BHPS asks women whether or not they expect to have a baby in the future. However, this question is not asked every year. Given my research design, I do not have enough data to use subsample by current fertility expectations. However, there is a strong correlation between expecting to have another child and actually having one.

not have more children, but the effects between the two groups are not significantly different. Mothers who do not have more children are significantly more likely to go back to work after the maternity leave ends than they would have been in the absence of job protection, and appear to stay more attached to the labor force as the child grows up. Column (6) indicates that Great Britain's maternity leave job protection expansion makes mothers 6 percentage points more likely to be employed five years after the birth of their last child. Because future fertility expectations should be unknown by employers, these results suggest that the effects of the maternity leave policies on employment are driven largely by mothers' choices in investments in work and childcare rather than employer discrimination. This idea is explored further below.

Education may be another important source of heterogeneity in the effects of wage replacement and job protection. Prior literature in the U.S. has shown that low-education mothers are more responsive to policies that provide wage replacement only, whereas more educated women might benefit more from job protection (Rossin-Slater et al., 2013; Rossin, 2011). However, Table 10 shows that there are no consistent differences in the employment patterns of high and low educated women under either policy reform. High educated women are defined as those with at least one year of of "further education" beyond the compulsory level. In the UK, compulsory schooling typically ends at age 16 at the completion of secondary school.[40] Unlike in the U.S., about 50 percent of mothers did not continue their education beyond secondary school. While the point estimates for the employment effects of job protection are in general higher for high educated mothers compared to low educated mothers, none are significantly different.

### 7.2 Effects on Job Characteristics

There are several mechanisms through which the maternity leave policies may affect the career trajectories of mothers that could explain the negative effects of job protection on earnings conditional on employment. First, job protection may change the types of jobs that mothers hold. Second, as shown above, it changes selection into employment.

By design, job protection should positively affect job tenure, defined as the number of months

[40]The GCSEs mark the end of the secondary education, and then individuals can either leave school or go on to further education in the form of either academic or vocational qualifications.

a worker has been in the same job. Column (1) of Table 11 shows that, conditional on employment, the job protection expansion increases average job tenure measured five years later by over 13 months, or 21 percent.[41] The eligibility expansion does not significantly affect long-run job tenure. Because the policies affect selection into employment at the extensive margin, column (2) shows the results hold when mothers who are not employed are included as well.

Appendix Table B4 shows that this increase in job tenure is accompanied by a large increase in the probability that a mother works for the same employer as she did before the child was born. Notably, the job protection reform increases the probability that a mother works for the same employer five years later by 13.4 percentage points. This represents a 31 percent increase in the probability of never changing employers among mothers who were employed in the year their child was born. These results are consistent with the large increase in average job tenure.

Although proponents of national maternity leave policies argue that maternity leave should help women rise to the top of the corporate ladder, there is little evidence of this in Great Britain. Despite the persistent increase in job tenure, mothers who are exposed to the maternity leave expansions are no more likely to be in a management or supervisory position or to have been promoted since the birth of their child than are other mothers. In fact, I find weak evidence in Column (3) of Table 11 that suggests the job protection expansion may actually make women less likely to become a manager or supervisor. This may be due to the fact that although job tenure increases, they are absent from work for longer periods than they would be in the absence of the policy. Column (4) shows that the eligibility expansion makes mothers over 6 percentage points (12 percent) less likely to have been promoted in the last five years. However, this result is entirely driven by mothers who have a subsequent birth. This supports the idea that women who anticipate having another child are working in order to qualify for additional wage replacement benefits in the future and are going into slightly worse than average jobs.[42]

There are several reasons why we might expect these intensive margin career effects to be larger in magnitude for highly educated women. For example, many high skill professionals

[41] I have defined this variable such that job tenure does not reset if a worker is promoted (and does not change employers), but job tenure resets to zero if the worker changes jobs within the firm and does not report being promoted.

[42] Appendix Table B5 shows the same outcomes for a sample of married women only. Again, the results are generally larger in magnitude for married women than for unmarried women.

work in occupations where human capital accumulation rates and promotion tracks are very steep and rigid during the early years of one's career. If individuals who work in these "up-or-out" environments miss a set window of opportunity for advancement, they are never able to do so (Rosen, 1990; Demougin and Siow, 1994; O'Flaherty and Siow, 1995). This window typically corresponds with prime childbearing years. Therefore, high skill mothers in these occupations may have a difficult time achieving high levels of professional success if they reduce short-run work investments or hours when they have children, even if the reduction is only temporary. Job protection, while encouraging women to stay employed and return to work soon after the period of leave ends, may not successfully mitigate the career penalty associated with the increased amount of leave taken.

Table 12 shows that the negative long-term effects of the job protection reform on the probability of being in a management position or having been promoted are concentrated among the more educated.[43] This implies that taking maternity leave, even while remaining officially employed, is costly for these women. This may be partially due to reduced human capital accumulation (or alternatively, increased skill depreciation). Although highly educated women are more likely to have longer firm tenure when job protection is available, the increased firm-specific human capital may not be enough to offset the cost of taking leave. Recent literature on returns to tenure suggests that, conditional on occupation tenure, the return to firm tenure is small (Kambourov and Manovskii, 2009). Using the BHPS, Zangelidis (2008) further finds substantial heterogeneity in the returns to firm tenure by skill. Ten years of firm tenure has no effect on the wages of workers in high-skill occupations including managers and professional occupations. The same amount of firm tenure has a large and significant positive effect on long-run wages for lower skill workers in clerical, service, sales, and craft occupations. Given I use the same data, it is not surprising that I also find no evidence of a return to increased firm tenure among these highly educated workers.

Finally, because the job protection reform increases the probability of being employed five years after the birth of a child, it is possible that the effects on becoming a manager or being

[43] As can be seen in Figure 4, the negative wage effects are concentrated at the top of the distribution, which is primarily comprised of high educated women.

promoted are driven by selection into the group of employed mothers. To determine how much of the effects are driven by selection, I perform a bounding exercise. Similar to Lee (2009) and Rossin-Slater (2016), I trim the sample of employed women by the number of "extra" women who select into employment as a result of the maternity leave policies.[44] The results from this bounding exercise are shown in Appendix Table B6 for the sample overall as well as by the mother's education level, and show that the negative effects I find on the probability of being a manager or having been promoted within 5 years cannot be explained by selection.

In my setting, the monotonicity assumption required for Lee bounds is not satisfied. For example, the job protection reform has a positive effect on long-run employment overall, likely driven by the "compliers" who return to work after a birth because of the presence of job protection. Monotonicity assumes that job protection only positively affects long-run employment decisions. But the policy likely induces some mothers who would have otherwise continued to work to instead take leave and then permanently exit the labor force. These "defiers" violate the monotonicity assumption. However, de Chaisemartin (2016) shows that the Lee bounds are still valid when monotonicity is violated as long as the percentage of response compliers is larger than the percentage of defiers and that the pre-treatment mean employment outcomes for the two groups are the same. Although not directly testable, these assumptions are both reasonable. The pre-reform share of employed mothers who want to give up paid work is small, suggesting that there are relatively few "defiers" compared to "compliers." Furthermore, the pre-reform long-term employment rates are the same for these two groups. It is therefore plausible that the bounds in Appendix Table B6 are valid.

### 7.3 Employer Hiring

Several economists propose that long maternity leave policies may lead to employer discrimination against women (Blau and Kahn, 2016). Mothers receive a lower return to work experience than women without children, even after controlling for time out of the labor market

[44] I implement this method by estimating the effect of each maternity leave reform on employment five years later separately by mother's education, as shown in Table 10. I then trim the group of employed mothers of infants in a particular policy/education group by that percent. To calculate the lower bound of the effect on each outcome, I drop only observations that have an outcome value of 1. To calculate the upper bound, I drop only observations that have a value of 0.

(Budig and Hodges, 2010; Waldfogel, 1998). Employers may penalize women for taking maternity leave or reducing their usual hours worked, even if it does not affect their long-term productivity (Bertrand et al., 2010). Anecdotal evidence also suggests that job-protected maternity leave may have unintended negative consequences for women of childbearing age, regardless of their motherhood status. While the direct costs to employers of providing wage replacement benefits to workers are minimal, the lost productivity costs associated with allowing women to go on job-protected leave may be substantial. In 2009, *Vogue* editor Alexandra Shulman published a now famous editorial in the UK tabloid *Daily Mail* entitled "Year-long maternity leave, flexi hours and four day weeks...why would ANY boss hire a woman?"[45] As the title suggests, Shulman posited that family-friendly policies including maternity leave in the UK may actually hurt women's employment prospects by raising the relative cost of hiring them. Others similarly suggest that job protected leave and other flexible hours policies are more costly for small employers who cannot afford to temporarily replace workers and do not have the personnel to easily distribute the extra work. It also seems likely that it is more difficult to temporarily replace high skill workers than low skill workers. If this is the case, then job protection might also be more costly for employers that hire more educated workers.

To explore the possibility that these policies are more costly for small employers, Table 13 looks at the DD effect of the maternity leave reforms on the probability that a woman works for a small (less than 50 employees) firm. In order to account for the fact that the policies might affect all women, these regressions compare the difference in outcomes of women before and after the policy changes to the same difference among men. Under the assumption that the gender-specific trends in employer size are not otherwise differentially changing during this period, the results suggest that small employers are less likely to hire female workers after the job protection reform. In these regressions, the treatment indicators are lagged three years to give employers time to learn about the costs of having women take leave and to potentially change their hiring practices. Job-protected leave appears to negatively affect the probability that women work at a small firm, and these results are driven by women of early childbearing age. This is what we

[45] *Daily Mail*, November 11, 2009, online edition: `http://www.dailymail.co.uk/debate/article-1226157/Vogue-editor-Alexandra-Shulman-asks-boss-hire-woman.html`. Accessed April 23, 2015.

would expect if employers are discriminating against women who are likely to start families in the near future. Columns (3) and (4) split the sample by education and show that the negative effects on hiring by small firms are entirely concentrated among more educated workers. This suggests that they are harder or more costly for small employers to replace during periods of leave.

But because mean monthly wages are about $£200$ lower on average for women who work in small firms compared to larger firms, it is not clear whether or not this is a bad outcome. If women have preferences for small firms (due to these firms offering better nonwage amenities, for example) or if this effect is driven by a decrease in the female employment rate, then the job protection policy may have a negative effect on young women overall.[46] But the negative effect of the job protection reform on the probability of working for a small firm is offset by an increase in the probability of working for a large firm. Table 14 shows that there is no effect on the overall female employment rate, nor are there differential effects by education. This policy also does not appear to have a differential impact on women's wages by firm size. Together, these two finding suggest it is possible that the overall effect on young women could be positive. While this analysis provides some evidence of job protection causing statistical discrimination against women at least within small firms, there is little reason to think it is driving the negative effects of job protection on measures of professional success.

There is also little evidence that the job protection reform is associated with changes in mothers' happiness with regard to work or life in general. Mothers do not report wanting to work a different number of hours after the policy goes into effect, change employers or jobs, or stop working altogether. There are also no differences in self-reported satisfaction with their relationship with their employer, income, or life in general. These findings should be viewed as descriptive rather than causal due to data limitations, but further suggest that the negative effects on hours worked and job characteristics could be driven by the work-related choices mothers are making rather than employer discrimination.[47]

[46] This effect does not appear to be driven by mothers in particular selecting into large firms, as both those with and without children are affected.

[47] Questions about preferences are not asked in early years of the survey, so I can only look at differences between the wage replacement and job protection reforms, and I do not find differences by education.These results are available upon request.

# 8 Conclusion

Maternity leave policies can have substantial effects on the employment outcomes and career trajectories of mothers, but the details of the policy matter. Using two large maternity leave policy reforms in Great Britain, I separate the effects of expanding access to maternity leave wage replacement benefits from the effects of additionally providing job protection benefits. I show that expanding access to paid leave has short-term effects on maternal employment, but these effects fade within three years. Furthermore, even these relatively short-run effects are concentrated among women who anticipate having more children in the future. This suggests that the increase in employment may be a result of women trying to qualify for additional periods of maternity leave benefits.

If policymakers want to increase women's employment rates in the long run, it appears that providing them with a substantial amount of job protection is important. My results suggest that Great Britain's job protection reform increases the probability that a mother is employed five years after the birth of her child by 8 percent, and these effects persist longer as well. It also increases average job tenure by over a year. However, there is no evidence that either type of maternity leave reform causes women to hold higher quality jobs or earn higher wages. This means that maternity leave may not be an effective way for policymakers to reduce the motherhood gap in earnings.

Although job protection increases firm tenure, it also increases the amount of leave that women take when their child is born. My results imply that the negative effect of the increased time off on long-term earnings and other measures of professional success outweigh any positive effects of increased job tenure, especially for high-skill women. Other recent work suggests that many high-skill careers are characterized by rigid promotion tracks and inflexible schedules that make it difficult to balance having both a family and successful career even when family-friendly benefits may be available (Antecol et al., 2016; Goldin and Katz, 2011, 2012). This indicates that it may be necessary to change the institutional structure of these careers in order to move closer toward gender equality in the workplace.

Finally, others argue that in order to achieve gender equality, family-friendly policies need

to promote equal participation of mothers and fathers in both childcare and market work (Gornick and Meyers, 2008). While job protection is an effective tool to increase job continuity and employment rates on the extensive margin, dedicated maternity leave induces only mothers to take more time off from work after the birth of a child. This may actually promote a gendered division of labor within the household. In fact, a 2009 report by the Equality and Human Rights Commission suggests that Great Britain's current maternity leave policy sends the message that women are primarily responsible for childrearing. It proposes decreasing the amount of dedicated maternity leave available and instead replacing it with parental leave that can be used by fathers as well (Equality and Human Rights Commission, 2009). In addition to encouraging equality in the home, higher male take-up of parental leave benefits should reduce employer statistical discrimination against women and necessitate institutional change in careers with rigid promotion structures that correspond with prime childbearing years.

Despite its shortcomings, maternity leave provides important benefits to mothers. In Great Britain, job protection clearly improves the long-term labor force attachment of mothers compared to wage replacement alone. This is an important finding given the large effect of work interruptions on future earnings, and should be carefully considered as the U.S. moves closer toward adopting a paid parental leave policy. Finally, it is important to consider that while taking maternity leave may be costly for mothers who intend to return to work, the time off may be very beneficial to infants (Stearns, 2015; Carneiro et al., 2015). Further research is needed to understand how wage replacement and job protection affect overall household welfare, as well as to determine the optimal length of each of these maternity leave benefits.

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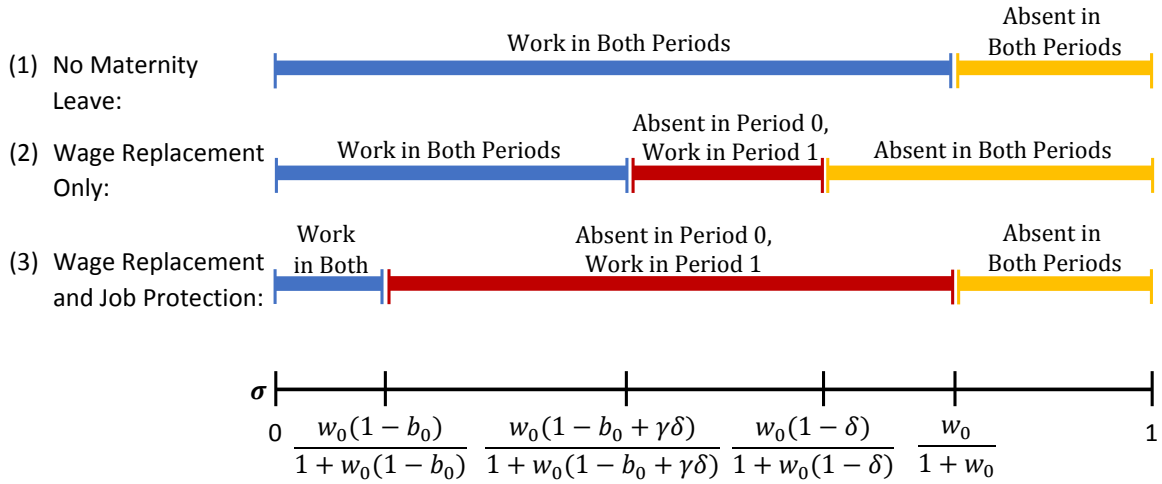
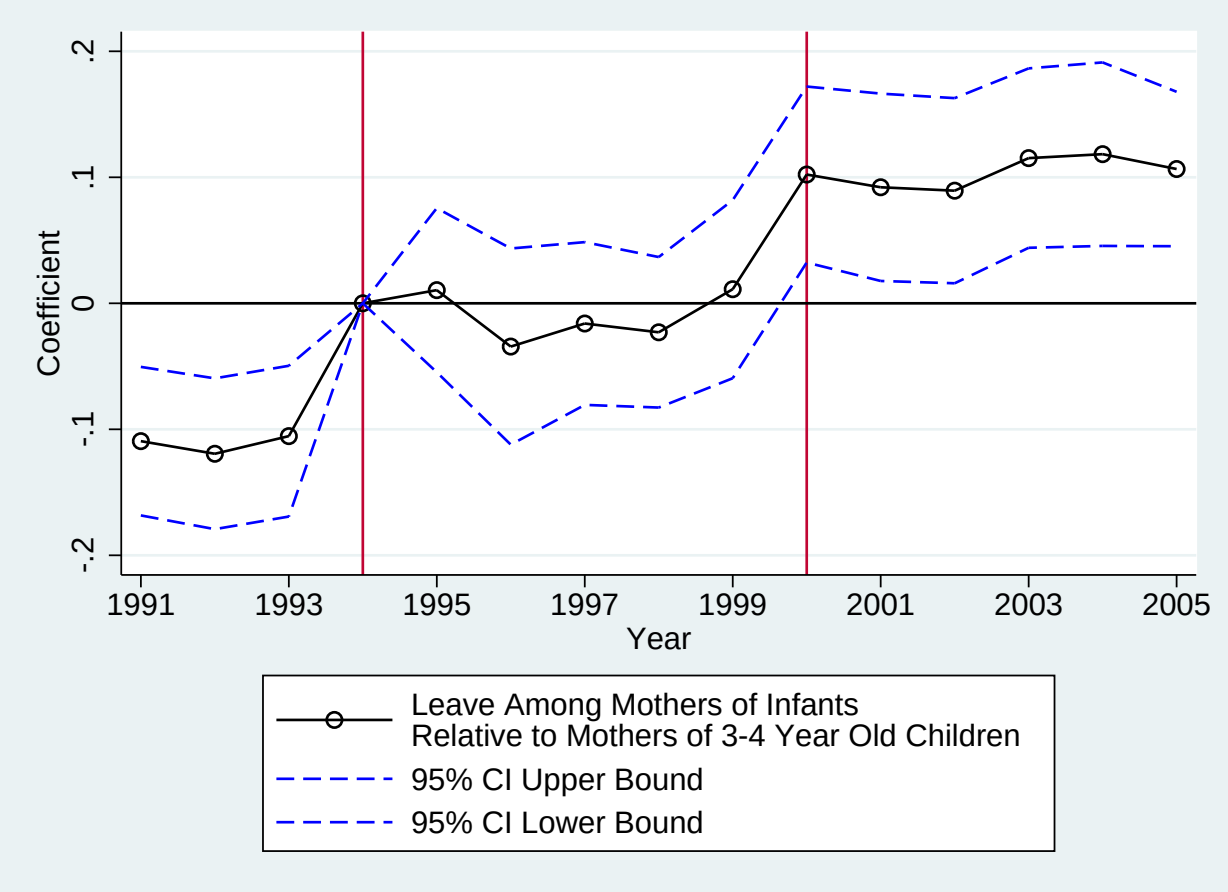
**Figure 1:** Optimal Work Decisions as the Relative Preference for Childrearing ($\sigma$) Varies



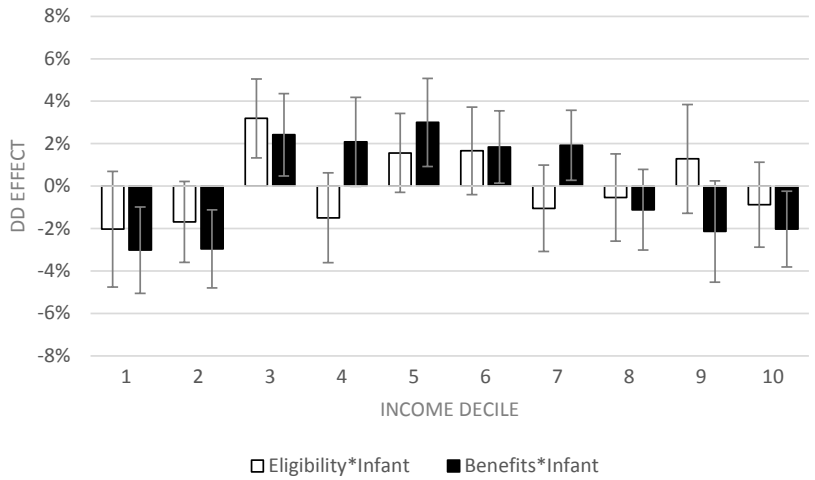

This figure shows the optimal work and childrearing decisions a mother will make in each period under no maternity leave policy, a maternity leave policy with wage replacement only, and a maternity leave policy with wage replacement and job protection. The $\sigma$ line at the bottom applies to all for cases above it. This ordering holds if $b_0 > \delta(1+\gamma)$.

**Figure 2:** Event Study Graph of Probability of Being on Leave; Mothers of Infants Compared to Mothers of 3-4 Year Olds



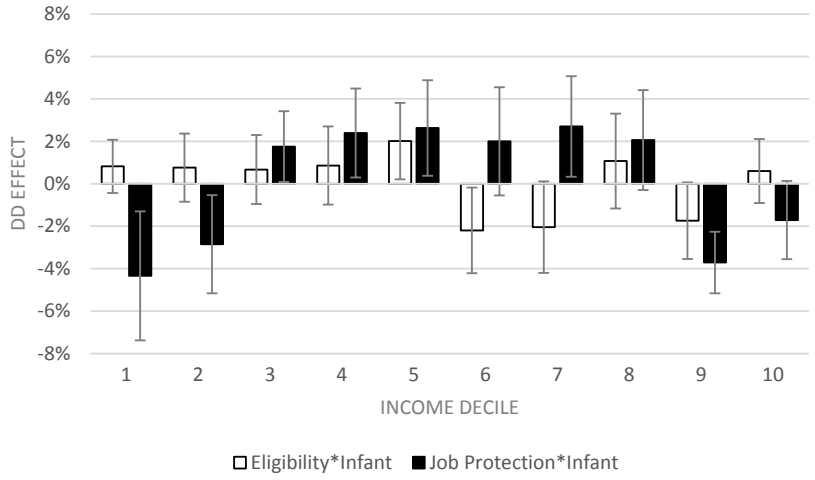

This figure plots the coefficients and 95 percent confidence intervals from an event-study regression that compares the leave-taking rate of mothers of infants relative to mothers whose youngest child is 3-4 years old in each year. The omitted year is 1994, and the vertical lines represent the introduction of the two maternity leave reforms.

**Figure 3:** Effect of Policies on Decile of Usual Monthly Income Distribution–Year of Birth




DD coefficients from ten separate regressions are shown, along with 95 percent confidence intervals. The dependent variable in each regression is an indicator for reporting an average monthly income within that decile. Regressions compare mothers of infants to those with a youngest child 3-4 in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

**Figure 4:** Effect of Policies on Decile of Usual Monthly Income Distribution–Five Years After Birth




DD coefficients from ten separate regressions are shown, along with 95 percent confidence intervals. The dependent variable in each regression is an indicator for reporting an average monthly income within that decile. Regressions compare mothers of infants to those with a youngest child 3-4 in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

**Table 1:** Summary Statistics

| | Whole Sample | Mothers of Child under 15 | Mothers of Infants | Mothers of Youngest Child Age 3-4 |
|---|---|---|---|---|
| Female | 0.530 | 1 | 1 | 1 |
| | (0.499) | (0) | (0) | (0) |
| Age | 47.238 | 37.636 | 29.328 | 33.189 |
| | (19.108) | (7.380) | (5.919) | (5.726) |
| Married | 0.548 | 0.690 | 0.660 | 0.680 |
| | (0.498) | (0.463) | (0.474) | (0.466) |
| Couple | 0.633 | 0.797 | 0.870 | 0.815 |
| | (0.482) | (0.402) | (0.336) | (0.388) |
| White | 0.959 | 0.946 | 0.937 | 0.943 |
| | (0.198) | (0.227) | (0.244) | (0.231) |
| Black | 0.011 | 0.017 | 0.015 | 0.017 |
| | (0.105) | (0.128) | (0.120) | (0.128) |
| Asian | 0.021 | 0.030 | 0.040 | 0.033 |
| | (0.142) | (0.171) | (0.195) | (0.178) |
| Other Race | 0.006 | 0.006 | 0.007 | 0.006 |
| | (0.080) | (0.078) | (0.082) | (0.078) |
| Number of Kids | 0.498 | 1.509 | 1.857 | 1.990 |
| | (0.923) | (1.039) | (1.049) | (0.923) |
| Infant | 0.027 | 0.084 | 1 | 0 |
| | (0.163) | (0.277) | (0) | (0) |
| Has Child | 0.653 | 1 | 1 | 1 |
| | (0.476) | (0) | (0) | (0) |
| Currently at Work | 0.561 | 0.641 | 0.330 | 0.629 |
| | (0.496) | (0.480) | (0.449) | (0.494) |
| Employed | 0.593 | 0.694 | 0.558 | 0.660 |
| | (0.491) | (0.461) | (0.500) | (0.488) |
| N | 225,601 | 44,005 | 4,024 | 7,233 |

Standard deviations in parentheses.

**Table 2:** Correlation between Maternity Leave Reforms and Mothers' Characteristics

| | (1) Has Infant | (2) Mean Age | (3) Married | (4) Couple | (5) Mean Age of Youngest Child | (6) High Education | (7) White | (8) Black | (9) Asian | (10) Other |
|---|---|---|---|---|---|---|---|---|---|---|
| Eligibility | 0.00503 | | | | | | | | | |
| | (0.00938) | | | | | | | | | |
| Job Protection | -0.00112 | | | | | | | | | |
| | (0.0105) | | | | | | | | | |
| Eligibility*Infant | | 0.0237 | -0.00448 | -0.0285 | -0.435 | -0.0157 | -0.0190 | -0.00473 | 0.0171 | 0.0001 |
| | | (0.0319) | (0.0445) | (0.0446) | (0.516) | (0.0429) | (0.0256) | (0.0164) | (0.0168) | (0.0116) |
| Job Protection*Infant | | 0.0136 | -0.00635 | -0.0314 | 0.291 | -0.0186 | 0.0233 | 0.0144 | -0.0320** | -0.00349 |
| | | (0.0212) | (0.0228) | (0.0212) | (0.300) | (0.0328) | (0.0181) | (0.00903) | (0.0147) | (0.00611) |
| Observations | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 | 11,257 |
| R-squared | 0.909 | 0.999 | 0.852 | 0.780 | 0.990 | 0.075 | 0.043 | 0.047 | 0.067 | 0.013 |
| Individual Controls | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | NO | YES | YES | YES | YES | NO | NO | NO | NO | NO |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables, excluding the outcome and other mutually exclusive dummy variables. High education is defined as completing any schooling beyond secondary school.

*** p<0.01, ** p<0.05, * p<0.1

**Table 3:** Effects of Maternity Leave on Short-Run Outcomes

| | (1) On Leave | (2) Currently Employed |
|---|---|---|
| Eligibility*Infant | 0.104*** | -0.0285 |
| | (0.0201) | (0.0282) |
| Job Protection*Infant | 0.102*** | 0.0377*** |
| | (0.0182) | (0.0156) |
| Observations | 11,257 | 11,257 |
| R-squared | 0.405 | 0.714 |
| Individual Controls | YES | YES |
| Time FE | YES | YES |
| Individual FE | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers of youngest children age 3-4. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.
*** p<0.01, ** p<0.05, * p<0.1

**Table 4:** Effects of Maternity Leave on Long-Run Employment

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | All Mothers | | | Married Mothers | | |
| | Employed in 1 year | Employed in 3 Years | Employed in 5 Years | Employed in 1 year | Employed in 3 Years | Employed in 5 Years |
| Eligibility*Infant | 0.0656** | 0.0368* | -0.0435 | 0.0797** | 0.0461* | 0.0292 |
| | (0.0323) | (0.0213) | (0.0579) | (0.0349) | (0.0249) | (0.0359) |
| Job Protection*Infant | 0.0209 | 0.0392* | 0.0454** | 0.0277 | 0.0397** | 0.0573** |
| | (0.0131) | (0.0233) | (0.0222) | (0.0229) | (0.0171) | (0.0230) |
| Observations | 9,927 | 9,743 | 9,247 | 6,491 | 5,112 | 5,993 |
| R-squared | 0.714 | 0.715 | 0.714 | 0.714 | 0.720 | 0.756 |
| Individual Controls | YES | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table 5:** Effects of Maternity Leave by Prior Eligibility

| | (1) Maternity Leave | (2) Employed This Year | (3) Employed in 1 Year | (4) Employed in 3 Years | (5) Employed in 5 Years |
|---|---|---|---|---|---|
| Eligibility*Infant | | | | | |
| Gained Eligibility | 0.146*** | -0.0606** | 0.0763 | 0.0477 | 0.0351 |
| | (0.0494) | (0.0279) | (0.0680) | (0.0605) | (0.0629) |
| Additional Effect for Already Eligible | -0.104*** | 0.0956 | -0.106 | -0.0329 | -0.0447 |
| | (0.0268) | (0.0926) | (0.0996) | (0.0618) | (0.122) |
| Job Protection*Infant | | | | | |
| Gained Eligibility Under | 0.111*** | 0.0715* | 0.0344 | 0.0694* | 0.0834** |
| Eligibility Reform | (0.0295) | (0.0370) | (0.0355) | (0.0360) | (0.0337) |
| Additional Effect for Already | -0.046 | -0.0252 | 0.00231 | -0.0134 | -0.0110 |
| Eligible Under Eligibility Reform | (0.0411) | (0.0585) | (0.0391) | (0.0363) | (0.0450) |
| Observations | 11,257 | 11,257 | 9,927 | 9,743 | 9,247 |
| R-squared | 0.527 | 0.716 | 0.716 | 0.719 | 0.764 |
| Individual Controls | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Gained Eligibility is an indicator for individuals who would not have been eligible for maternity leave under the pre-1994 policy, measured by their employment history in 1993. Already Eligible is an indicator for individuals who would have been eligible in 1993. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

**Table 6:** Effect of Maternity Leave on Usual Hours Worked

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) |
|---|---|---|---|---|---|---|---|---|
| | Usual Hours Per Week (Conditional on Working) | | | | Usual Hours Per Week (Unconditional) | | | |
| | This Year | in 1 Year | in 3 Years | in 5 Years | This Year | in 1 Year | in 3 Years | in 5 Years |
| Eligibility*Infant | -0.444 | -1.686 | 0.424 | -0.553 | -0.195 | 2.625 | 0.888 | 0.453 |
| | (1.594) | (1.526) | (1.467) | (1.323) | (1.355) | (1.289) | (1.422) | (1.451) |
| Job Protection*Infant | -2.637*** | -3.175*** | -0.695 | -0.490 | 0.574 | -2.210*** | 1.370* | 1.336 |
| | (0.901) | (0.879) | (0.970) | (1.027) | (0.787) | (0.712) | (0.830) | (0.917) |
| Observations | 5,670 | 5,354 | 4,785 | 4,174 | 10,669 | 9,927 | 9,743 | 9,247 |
| R-squared | 0.113 | 0.082 | 0.052 | 0.053 | 0.752 | 0.767 | 0.757 | 0.789 |
| Individual Controls | YES | YES | YES | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables. Columns (1)-(4) limit the sample to employed mothers; columns (5)-(8) do not condition on employment.

*** p<0.01, ** p<0.05, * p<0.1

**Table 7:** Effect of Maternity Leave on Probability of Working Full-Time

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | Full-time Work (>35 Hours Per Week) | | | |
| | This Year | in 1 Year | in 3 Years | in 5 Years |
| Eligibility*Infant | 0.00222 | -0.0514 | -0.00591 | -0.0353 |
| | (0.0611) | (0.0578) | (0.0524) | (0.0500) |
| Job Protection*Infant | -0.0684* | -0.133*** | -0.0571 | -0.0544 |
| | (0.0372) | (0.0350) | (0.0372) | (0.0397) |
| Observations | 5,670 | 5,354 | 4,785 | 4,174 |
| R-squared | 0.096 | 0.059 | 0.032 | 0.042 |
| Individual Controls | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES |
| Individual FE | NO | NO | NO | NO |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Sample is conditional on being employed. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table 8:** Effect of Maternity Leave on Long-Run Mean Monthly Income

| | (1) | (2) |
|---|---|---|
| | Mean Earnings in 5 Years | |
| | All Women | Conditional on Employment |
| Eligibility*Infant | 48.39 | -42.48 |
| | (50.35) | (68.59) |
| Job Protection*Infant | 10.22 | -70.90* |
| | (33.23) | (37.47) |
| Mean of Outcome | £972 | £1,304 |
| Observations | 9,247 | 6,202 |
| R-squared | 0.695 | 0.791 |
| Individual Controls | YES | YES |
| Time FE | YES | YES |
| Individual FE | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table 9:** Effects of Maternity Leave on Long-Run Employment, by Future Fertility

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Mother Has More Children | | | No More Children | | |
| | Employed in 1 year | Employed in 3 Years | Employed in 5 Years | Employed in 1 year | Employed in 3 Years | Employed in 5 Years |
| Eligibility*Infant | 0.0686*** | 0.0433*** | -0.0127 | -0.0311 | -0.0233 | -0.0222 |
| | (0.0213) | (0.0102) | (0.0160) | (0.0369) | (0.0244) | (0.0376) |
| Job Protection*Infant | 0.0399 | 0.0340 | 0.0418 | 0.0650* | 0.0607*** | 0.0644** |
| | (0.0331) | (0.0362) | (0.0355) | (0.0346) | (0.0304) | (0.0317) |
| Observations | 4,062 | 4,062 | 4,062 | 5,117 | 5,117 | 5,117 |
| R-squared | 0.694 | 0.784 | 0.857 | 0.763 | 0.760 | 0.801 |
| Individual Controls | YES | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Columns (1)-(3) condition the sample on women who have at least one birth after the reference year. Columns (4)-(6) condition the sample on women who do not have more children after the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

**Table 10:** Effects of Maternity Leave on Employment by Mother's Education

| | (1) Employed This Year | (2) Employed in 1 Year | (3) Employed in 3 Years | (4) Employed in 5 Years |
|---|---|---|---|---|
| *Panel A: Mothers with High Education* | | | | |
| Eligibility*Infant | -0.00707 | 0.00214 | 0.0125 | 0.0316 |
| | (0.0513) | (0.0822) | (0.0872) | (0.0876) |
| Job Protection*Infant | 0.0617** | 0.0252 | 0.0513* | 0.0419* |
| | (0.0283) | (0.0290) | (0.0300) | (0.0222) |
| Observations | 4,381 | 4,132 | 3,576 | 3,507 |
| R-squared | 0.720 | 0.723 | 0.734 | 0.759 |
| *Panel B: Mothers with Low Education* | | | | |
| Eligibility*Infant | -0.0277 | 0.0770* | 0.0445 | 0.0124 |
| | (0.0710) | (0.0448) | (0.0453) | (0.0738) |
| Job Protection*Infant | 0.0379*** | 0.0103 | 0.0247 | 0.0462* |
| | (0.0183) | (0.0284) | (0.0295) | (0.0267) |
| Observations | 4,689 | 4,452 | 3,987 | 3,450 |
| R-squared | 0.699 | 0.709 | 0.714 | 0.738 |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. High education is defined as one year of post-secondary education or more. Low education is defined as completing secondary education or less. Regressions include time and individual fixed effects, and individual controls including age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

**Table 11:** Effects of Maternity Leave on Long-Run Job Characteristics

| | (1) Job Tenure (Months) in 5 Years | (2) Job Tenure (Months) Including Not Employed in 5 Years | (3) Is a Manager in 5 Years | (4) Promoted within 5 Years | (5) Promoted within 5 Years No More Kids |
|---|---|---|---|---|---|
| Eligibility*Infant | 6.133 | 2.84 | -0.0943 | -0.0657** | 0.0085 |
| | (6.565) | (3.724) | (0.0618) | (0.0305) | (0.0829) |
| Job Protection*Infant | 13.742** | 7.664** | -0.0646* | 0.0213 | -0.0315 |
| | (6.268) | (3.311) | (0.0386) | (0.0399) | (0.0613) |
| Observations | 4,115 | 6,123 | 3,625 | 3,454 | 2,254 |
| R-squared | 0.791 | 0.795 | 0.798 | 0.648 | 0.717 |
| Individual Controls | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table 12:** Effects of Maternity Leave on Long-Run Job Characteristics by Education

| | (1) Job Tenure (Months) in 5 Years | (2) Job Tenure (Months) Including Not Employed in 5 Years | (3) Is a Manager in 5 Years | (4) Promoted within 5 Years |
|---|---|---|---|---|
| *Panel A: Mothers with High Education* | | | | |
| Eligibility*Infant | 1.198 | 3.658 | -0.0518 | 0.00267 |
| | (12.26) | (7.456) | (0.223) | (0.0805) |
| Job Protection*Infant | 19.77** | 8.136** | -0.125*** | -0.0749* |
| | (8.082) | (4.001) | (0.0521) | (0.0402) |
| Mean of Outcome | 68.26 | 0.444 | 0.484 | |
| Observations | 1,593 | 2,810 | 1,634 | 1,576 |
| R-squared | 0.768 | 0.789 | 0.797 | 0.710 |
| *Panel B: Mothers with Low Education* | | | | |
| Eligibility*Infant | 5.822 | 1.550 | -0.0509 | -0.106*** |
| | (8.871) | (4.854) | (0.152) | (0.0451) |
| Job Protection*Infant | 6.093 | 6.762 | 0.0637 | 0.0633 |
| | (7.440) | (4.171) | (0.111) | (0.0917) |
| Mean of Outcome | 44.13 | 0.258 | 0.445 | |
| Observations | 1,522 | 3,313 | 1,577 | 1,502 |
| R-squared | 0.835 | 0.800 | 0.747 | 0.704 |

Cluster-robust standard errors in parentheses. Each column in each panel is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Regressions include time and individual fixed effects, and individual controls including age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

**Table 13:** Effects of Maternity Leave on Firm Size

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | Outcome: Employed by Small Firm 3 Years after Policy | | | |
| | All Workers | 21-34 Year Olds | High Education 21-34 Year Olds | Low Education 21-34 Year Olds |
| Eligibility*Female | -0.00426 | -0.00434 | -0.0415 | 0.0407 |
| | (0.0120) | (0.0268) | (0.0339) | (0.0347) |
| Job Protection*Female | -0.0339*** | -0.0607*** | -0.0926*** | 0.00190 |
| | (0.0080) | (0.0217) | (0.0363) | (0.0277) |
| Mean of Outcome | 0.475 | 0.463 | 0.445 | 0.473 |
| Observations | 52,679 | 14,770 | 5,978 | 8,792 |
| R-squared | 0.643 | 0.696 | 0.697 | 0.672 |
| Individual Controls | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing females to males. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables. A small firm is defined as having less than 50 workers. High education is defined as one year of post-secondary education or more. Low education is defined as completing secondary education or less.

*** p<0.01, ** p<0.05, * p<0.1

**Table 14:** Effects of Maternity Leave on Overall Female Employment

| | (1) | (2) | (3) | (4) |
|---|---|---|---|---|
| | Outcome: Employed 3 Years after Policy | | | |
| | | | High Education | Low Education |
| | All Workers | 21-34 Year Olds | 21-34 Year Olds | 21-34 Year Olds |
| Eligibility*Female | 0.0119 | 0.00008 | 0.0118 | -0.0150 |
| | (0.00894) | (0.0181) | (0.0280) | (0.0238) |
| Job Protection*Female | 0.00663 | 0.00253 | 0.000933 | 0.00493 |
| | (0.00600) | (0.0162) | (0.0275) | (0.0201) |
| Mean of Outcome | 0.790 | 0.771 | 0.804 | 0.751 |
| Observations | 84,941 | 25,990 | 11,146 | 14,844 |
| R-squared | 0.597 | 0.658 | 0.688 | 0.64 |
| Individual Controls | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing females to males. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables. High education is defined as one year of post-secondary education or more. Low education is defined as completing secondary education or less.

*** p<0.01, ** p<0.05, * p<0.1

# Appendix A Theoretical Model Details

## A.1 No Maternity Leave

When $b_0 = 0$,

$$
\begin{aligned}
U(W_0W_1) &= (1-\sigma)[I_0 + \gamma I_1 + w_0 + \gamma w_0] \\
U(K_0W_1) &= (1-\sigma)[I_0 + \gamma I_1 + \gamma w_0 - \gamma\delta w_0] + \sigma \\
U(W_0K_1) &= (1-\sigma)[I_0 + \gamma I_1 + w_0] + \sigma\gamma \\
U(K_0K_1) &= (1-\sigma)[I_0 + \gamma I_1] + \sigma + \sigma\gamma
\end{aligned}
$$

Then the following conditions must hold:

$$
\begin{aligned}
U(W_0W_1) > U(K_0W_1) \text{ if } \sigma &< \frac{w_0(1+\gamma\delta)}{1+w_0(1+\gamma\delta)} \\
U(W_0W_1) > U(W_0K_1) \text{ if } \sigma &< \frac{w_0}{1+w_0} \\
U(W_0W_1) > U(K_0K_1) \text{ if } \sigma &< \frac{w_0}{1+w_0} \\
U(K_0W_1) > U(W_0K_1) \text{ if } \sigma &< \frac{w_0(-1+\gamma(1-\delta))}{-1+\gamma+w_0(-1+\gamma(1-\delta))} \\
U(K_0W_1) > U(K_0K_1) \text{ if } \sigma &< \frac{w_0(1-\delta)}{1+w_0(1-\delta)} \\
U(W_0K_1) > U(K_0K_1) \text{ if } \sigma &< \frac{w_0}{1+w_0}
\end{aligned}
$$

Because $\delta > 0$ and $\gamma > 0$,

$$
\frac{w_0(1+\gamma\delta)}{1+w_0(1+\gamma\delta)} > \frac{w_0}{1+w_0} > \frac{w_0(1-\delta)}{1+w_0(1-\delta)} > \frac{w_0(-1+\gamma(1-\delta))}{-1+\gamma+w_0(-1+\gamma(1-\delta))}
$$

Therefore, $\sigma^*_{N0} = \sigma^*_{N1} = \frac{w_0}{1+w_0}$. If $\sigma < \frac{w_0}{1+w_0}$ then the mother will work in both periods; otherwise she will engage in childrearing in both periods.

## A.2 Maternity Leave with Wage Replacement

When $b_0 > 0$, $0 < \delta < 1$, and $0 < \gamma < 1$,

$$
\begin{aligned}
U(W_0W_1) &= (1-\sigma)[I_0 + \gamma I_1 + w_0 + \gamma w_0] \\
U(K_0W_1) &= (1-\sigma)[I_0 + \gamma I_1 + b_0 w_0 + \gamma w_0 - \gamma\delta w_0] + \sigma \\
U(W_0K_1) &= (1-\sigma)[I_0 + \gamma I_1 + w_0] + \sigma\gamma \\
U(K_0K_1) &= (1-\sigma)[I_0 + \gamma I_1 + b_0 w_0] + \sigma + \sigma\gamma
\end{aligned}
$$

Conditional on working in period 0, potential wage is the same in period 1 as in period 0 but the potential benefit to childrearing is strictly lower. Therefore, it is never optimal to choose to work in period 0 but not work in period 1. As long as $\gamma > 0$, mothers maximize lifetime utility by choosing between the three strategies $(W_0W_1)$, $(K_0W_1)$, and $(K_0K_1)$.

Then there are two possible cases. First, if $b_0 > \gamma\delta$ and $b_0 > \delta(1+\gamma)$,

$$
\begin{aligned}
&U(W_0W_1) > U(K_0W_1) > U(K_0K_1) \text{ if } \sigma < \frac{w_0(1-b_0+\gamma\delta)}{1+w_0(1-b_0+\gamma\delta)} \\
&U(K_0W_1) > U(W_0W_1) > U(K_0K_1) \text{ if } \frac{w_0(1-b_0+\gamma\delta)}{1+w_0(1-b_0+\gamma\delta)} \leq \sigma \leq \frac{w_0(1-\delta)}{1+w_0(1-\delta)} \\
&U(K_0K_1) > U(K_0W_1) > U(W_0W_1) \text{ if } \sigma > \frac{w_0(1-\delta)}{1+w_0(1-\delta)}
\end{aligned}
$$

Therefore, if $\sigma < \frac{w_0(1-b_0+\gamma\delta)}{1+w_0(1-b_0+\gamma\delta)}$ the mother will work in both periods; if $\frac{w_0(1-b_0+\gamma\delta)}{1+w_0(1-b_0+\gamma\delta)} \leq \sigma \leq \frac{w_0(1-\delta)}{1+w_0(1-\delta)}$ she will work only in the later period; and if $\sigma > \frac{w_0(1-\delta)}{1+w_0(1-\delta)}$ she will not work in either period.

If $b_0 < \delta(1+\gamma)$, then the mother will always make the same labor supply decision in both periods. If $\sigma < \frac{w_0(1-b_0+\gamma)}{1+\gamma+w_0(1-b_0+\gamma)}$, she will work in both periods; otherwise she will engage in childrearing in both periods.

### A.3 Maternity Leave with Wage Replacement and Job Protection

When $\delta = 0$,

$$\sigma_J^{*0} = \frac{w_0(1-b_0)}{1+w_0(1-b_0)}$$
$$\sigma_J^{*1} = \frac{w_0}{1+w_0}$$

The mother will work if $\sigma < \sigma_J^{*t}$ and engage in childrearing otherwise. Because

$$\frac{w_0(1-b_0+\gamma)}{1+w_0(1-b_0+\gamma)} < \frac{w_0}{1+w_0}$$

the mother will work in both periods when $\sigma < \sigma_J^{*0}$, will stay home in the period 0 and work in period 1 when $\sigma_J^{*0} < \sigma < \sigma_J^{*1}$, and will stay home in both periods when $\sigma > \sigma_J^{*1}$.

### A.4 Allowing $\sigma_t$ to be Determined at the Start of Each Period

It may be more realistic to assume that mothers differentially value childrearing as their children age. It is possible to generalize the model to allow for $\sigma_t$ to vary across periods, and for this parameter to be realized at the start of each period. For simplicity, suppose $\sigma_t$ is drawn from a uniform distribution.

An individual will be indifferent between working and not working in period 0 when the utility from working in period 0 plus period 1 utility conditional on working in period 0 is equal to the utility from not working in period 0 plus period 1 utility conditional on not working in period 0:

$$\begin{aligned}(1-\sigma_0^*)w_0 + \gamma[\sigma_1^{W_0^*}(1-\frac{\sigma_1^{W_0^*}}{2})w_0 + (1-\sigma_1^{W_0^*})(\frac{1+\sigma_1^{W_0^*}}{2})] \\ = (1-\sigma_0^*)b_0 w_0 + \sigma_0^* + \gamma[\sigma_1^{K_0^*}(1-\frac{\sigma_1^{K_0^*}}{2})w_0(1-\delta) + (1-\sigma_1^{K_0^*})(\frac{1+\sigma_1^{K_0^*}}{2})]\end{aligned}$$

where $\sigma_1^{W_0^*} = \frac{w_0}{1-w_0}$ is the indifference point for working in period 1 conditional on working in period 0, and $\sigma_1^{K_0^*} = \frac{w_0(1-\delta)}{1-w_0(1-\delta)}$ is the indifference point conditional on childrearing in period 0. The expected period 1 utilities are a weighted average of the utility from choosing work and chil-

drearing, where the weights on working and childrearing are the probability that an individual's realized value of $\sigma_1$ will be below or above the relevant indifference point, respectively.

Solving for $\sigma_0^*$ yields:

$$\sigma_0^*(1+w_0(1-b_0)) = w_0(1-b_0) + \frac{\gamma}{2}[\frac{w_0(w_0^2+w_0-1)}{(1+w_0)^2}) - \frac{w_0(1-\delta)(w_0^2(1-\delta)^2+w_0(1-\delta)-1)}{(1+w_0(1-\delta))^2})]$$

$$\sigma_0^* = \frac{w_0(1-b_0)}{1+w_0(1-b_0)} + \frac{\gamma}{2}[\frac{w_0(w_0^2+w_0-1)}{(1+w_0(1-b_0))(1+w_0)^2} - \frac{w_0(1-\delta)(w_0^2(1-\delta)^2+w_0(1-\delta)-1)}{(1+w_0(1-b_0))(1+w_0(1-\delta))^2}]$$

It is clear that this indifferent point depends on both the maternity leave wage replacement rate, $b_0$, and the wage penalty for not working, $\delta$. By taking the derivative of $\sigma_0^*$ with respect to $b_0$, it is also evident that increasing maternity leave benefits affects labor supply in period 0 through both the direct effect in period 0 as well as the expectation about the wage loss in period 1.

$$\frac{\partial \sigma_0^*}{\partial b_0} = \frac{-w_0}{(1+w_0(1-b_0))^2} + \frac{\gamma}{2}[\frac{w_0^2(w_0^2+w_0-1)}{(1+w_0(1-b_0))^2(1+w_0)^2} - \frac{w_0^2(1-\delta)^2(w_0^2(1-\delta)^2+w_0(1-\delta)-1)}{(1+w_0(1-b_0))^2(1+w_0(1-\delta))^2}]$$

The first term is negative and represents the effect of increasing maternity leave benefits in a one period model, where childrearing in period 0 has no effect on future employment decisions. This captures the change in the income gap between working and childrearing in period 0. As $b_0$ increases, this gap diminishes and shifts the indifference point to the left. This means that more mothers choose childrearing in period 0. The next term is positive and therefore mitigates the direct effect. This term captures the fact that if increasing $b_0$ lowers the probability of working in period 0, it also lowers the utility of working in period 1. This penalty effectively lowers the value of $b_0$ for women who would go back to work in the next period, and shifts the indifference point back to the right relative to the case where there is no wage penalty for not working in period 0. If $\delta = 0$, the second term is 0.

In period 1, the mother chooses whether or not to work given her decision in period 0 and her realization of $\sigma_1$. Therefore, the indifference point $\sigma_1^*$ is the same as described above for the case where $\sigma$ is constant over time.

# Appendix B Additional Tables

**Table B1:** Effects of Maternity Leave on Leave-Taking Behavior: Alternate Specifications

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Control Group: | | | | | | | | |
| | Women with Youngest Child Age 1-2 | | | Mothers with Youngest Child Age 3-4 | | | Mothers with Youngest Child Under Age 15 | | |
| Panel A | | | | | | | | | |
| Eligibility*Infant | 0.0909*** | 0.0925*** | 0.110*** | 0.0911*** | 0.0897*** | 0.104*** | 0.0914*** | 0.0931*** | 0.101*** |
| | (0.0196) | (0.0191) | (0.0220) | (0.0194) | (0.0191) | (0.0201) | (0.0194) | (0.0192) | (0.0198) |
| Job Protection*Infant | 0.110*** | 0.110*** | 0.105*** | 0.105*** | 0.111*** | 0.102*** | 0.109*** | 0.107*** | 0.0976*** |
| | (0.0178) | (0.0172) | (0.0184) | (0.0178) | (0.0173) | (0.0182) | (0.0177) | (0.0173) | (0.0178) |
| Observations | 10,669 | 10,669 | 10,669 | 11,257 | 11,257 | 11,257 | 43,891 | 43,891 | 43,891 |
| R-squared | 0.171 | 0.209 | 0.439 | 0.166 | 0.207 | 0.405 | 0.217 | 0.237 | 0.395 |
| Individual Controls | NO | YES | YES | NO | YES | YES | NO | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | NO | NO | YES | NO | NO | YES | NO | NO | YES |

| | | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|
| | Control Group: | | | | | | | | |
| | Women without Children | | | Fathers with Infant | | | DDD (Fathers of Infants and Parents of Youngest Child Age 3-4) | | |
| Panel B | | | | | | | | | |
| Eligibility*Infant | 0.0933*** | 0.0981*** | 0.146*** | | | | | | |
| | (0.0195) | (0.0190) | (0.0305) | | | | | | |
| Job Protection*Infant | 0.109*** | 0.104*** | 0.105*** | | | | | | |
| | (0.0177) | (0.0171) | (0.0231) | | | | | | |
| Eligibility*Female | | | | 0.0918*** | 0.0905*** | 0.0721* | | | |
| | | | | (0.0194) | (0.0193) | (0.0406) | | | |
| Job Protection*Female | | | | 0.106*** | 0.104*** | 0.0927*** | | | |
| | | | | (0.0177) | (0.0174) | (0.0122) | | | |
| Eligibility*Female*Infant | | | | | | | 0.0903*** | 0.0901*** | 0.105*** |
| | | | | | | | (0.0197) | (0.0195) | (0.0213) |
| Job Protection*Female*Infant | | | | | | | 0.106*** | 0.106*** | 0.108*** |
| | | | | | | | (0.0179) | (0.0176) | (0.0185) |
| Observations | 32,891 | 32,891 | 32,891 | 7,308 | 7,308 | 7,308 | 19,231 | 19,231 | 19,231 |
| R-squared | 0.206 | 0.240 | 0.483 | 0.145 | 0.186 | 0.721 | 0.204 | 0.224 | 0.451 |
| Individual Controls | NO | YES | YES | NO | YES | YES | NO | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | NO | NO | YES | NO | NO | YES | NO | NO | YES |

Cluster-robust standard errors in parentheses. Each column in each panel is a separate regression comparing mothers of infants to the control group. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table B2:** Effects of Maternity Leave on the Probability of Being Currently Employed: Alternate Specifications

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Control Group: | | | | | | | | |
| | Mothers with Youngest Child Age 1-2 | | | Mothers with Youngest Child Age 3-4 | | | Mothers with Youngest Child Under Age 15 | | |
| Panel A | | | | | | | | | |
| Eligibility*Infant | 0.0114 | 0.0158 | -0.00748 | 0.00515 | -0.0112 | -0.0285 | 0.0391 | 0.0478* | 0.0305 |
| | (0.0382) | (0.0355) | (0.0363) | (0.0407) | (0.0386) | (0.0582) | (0.0312) | (0.0290) | (0.0533) |
| Job Protection*Infant | 0.0640** | 0.0707*** | 0.0454*** | 0.0537* | 0.0699*** | 0.0377*** | 0.0847*** | 0.0809*** | 0.0395** |
| | (0.0271) | (0.0245) | (0.0185) | (0.0281) | (0.0259) | (0.0156) | (0.0222) | (0.0205) | (0.0190) |
| Observations | 10,669 | 10,669 | 10,669 | 11,257 | 11,257 | 11,257 | 43,891 | 43,891 | 43,891 |
| R-squared | 0.018 | 0.168 | 0.705 | 0.028 | 0.148 | 0.714 | 0.020 | 0.115 | 0.630 |
| Individual Controls | NO | YES | YES | NO | YES | YES | NO | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | NO | NO | YES | NO | NO | YES | NO | NO | YES |
| | Control Group: | | | | | | | | |
| | Women without Children | | | Fathers with Infant | | | DDD (Fathers of Infants and Parents of Youngest Child Age 3-4) | | |
| Panel B | | | | | | | | | |
| Eligibility*Infant | 0.0999*** | 0.0948*** | 0.0569 | | | | | | |
| | (0.0314) | (0.0295) | (0.0373) | | | | | | |
| Job Protection*Infant | 0.122*** | 0.104*** | 0.0699*** | | | | | | |
| | (0.0223) | (0.0207) | (0.0237) | | | | | | |
| Eligibility*Female | | | | 0.0479 | 0.0478 | 0.0335 | | | |
| | | | | (0.0395) | (0.0366) | (0.0661) | | | |
| Job Protection*Female | | | | 0.0614** | 0.0564** | 0.0241 | | | |
| | | | | (0.0261) | (0.0243) | (0.0525) | | | |
| Eligibility*Female*Infant | | | | | | | -0.00872 | -0.00807 | -0.0166 |
| | | | | | | | (0.0502) | (0.0468) | (0.0408) |
| Job Protection*Female*Infant | | | | | | | 0.0572* | 0.0616** | 0.0588** |
| | | | | | | | (0.0331) | (0.0309) | (0.0241) |
| Observations | 32,891 | 32,891 | 32,891 | 7,308 | 7,308 | 7,308 | 19,231 | 19,231 | 19,231 |
| R-squared | 0.054 | 0.119 | 0.563 | 0.188 | 0.299 | 0.853 | 0.169 | 0.265 | 0.753 |
| Individual Controls | NO | YES | YES | NO | YES | YES | NO | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | NO | NO | YES | NO | NO | YES | NO | NO | YES |

Cluster-robust standard errors in parentheses. Each column in each panel is a separate regression comparing mothers of infants to the control group. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table B3:** Effects of Maternity Leave on Long-Run Employment: Alternate Specifications

| | (1) | (2) | (3) | (4) | (5) | (6) | (7) | (8) | (9) |
|---|---|---|---|---|---|---|---|---|---|
| | Control Group: | | | | | | | | |
| | Women without Children | | | Women with Youngest Child Age 3-15 | | | Fathers of Infants | | |
| | Employed in 1 year | Employed in 3 Years | Employed in 5 Years | Employed in 1 year | Employed in 3 Years | Employed in 5 Years | Employed in 1 year | Employed in 3 Years | Employed in 5 Years |
| Eligibility*Infant | 0.0650** | 0.0451 | 0.0337 | 0.0502*** | 0.0409 | 0.0325 | 0.0733 | 0.0366 | 0.0168 |
| | (0.0323) | (0.0303) | (0.0305) | (0.0195) | (0.0307) | (0.0294) | (0.0669) | (0.0730) | (0.0680) |
| Job Protection*Infant | 0.0445* | 0.0411* | 0.0624** | 0.0121 | 0.0615*** | 0.0449** | 0.0450 | 0.0399 | 0.0489* |
| | (0.0234) | (0.0235) | (0.0256) | (0.0186) | (0.0194) | (0.0196) | (0.0293) | (0.0261) | (0.0291) |
| Observations | 30,999 | 26,943 | 22,739 | 41,489 | 36,392 | 30,926 | 6,919 | 6,076 | 5,153 |
| R-squared | 0.530 | 0.612 | 0.700 | 0.608 | 0.626 | 0.695 | 0.845 | 0.851 | 0.867 |
| Individual Controls | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to the control group in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table B4:** Effects of Maternity Leave on the Probability of Working for the Same Employer

| | (1) Same Employer as Last Year | (2) Same Employer as Last Year (All Individuals) | (3) Same Employer in 5 Years | (4) Same Employer in 5 Years (All Individuals) |
|---|---|---|---|---|
| Eligibility*Infant | -0.0612 | -0.00235 | -0.0213 | 0.00734 |
| | (0.0937) | (0.0470) | (0.0971) | (0.0406) |
| Job Protection*Infant | 0.162*** | 0.108*** | 0.134*** | 0.0854*** |
| | (0.0496) | (0.0298) | (0.0507) | (0.0257) |
| Mean of Dependent Variable | 0.620 | 0.260 | 0.427 | 0.179 |
| Observations | 5,202 | 9,283 | 5,202 | 9,283 |
| R-squared | 0.597 | 0.663 | 0.718 | 0.651 |
| Individual Controls | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables. Same employer is an indicator for working for the same employer as last year and same employer in 5 years is an indicator for working for the same employer as 5 years ago. Columns (1) and (3) condition the sample on mothers who were employed in the year of birth; columns (2) and (4) do not.

*** p<0.01, ** p<0.05, * p<0.1

**Table B5:** Effects of Maternity Leave on Long-Run Job Characteristics: Married Mothers

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| | | | Married Women | | |
| | Job Tenure (Months) in 5 Years | Job Tenure (Months) Including Not Employed in 5 Years | Is a Manager in 5 Years | Promoted within 5 Years | Promoted within 5 Years No More Kids |
| Eligibility*Infant | 3.219 | 0.795 | 0.0259 | -0.173*** | 0.0283 |
| | (7.434) | (4.668) | (0.0655) | (0.0652) | (0.0863) |
| Job Protection*Infant | 17.034** | 10.152** | -0.115** | 0.0729 | -0.00926 |
| | (7.028) | (4.479) | (0.0465) | (0.0479) | (0.0707) |
| Observations | 2,296 | 4,054 | 2,343 | 2,566 | 1,164 |
| R-squared | 0.792 | 0.796 | 0.801 | 0.656 | 0.729 |
| Individual Controls | YES | YES | YES | YES | YES |
| Time FE | YES | YES | YES | YES | YES |
| Individual FE | YES | YES | YES | YES | YES |

Cluster-robust standard errors in parentheses. Each column is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and education dummy variables.

*** p<0.01, ** p<0.05, * p<0.1

**Table B6:** Effects of Maternity Leave on Long-Run Job Characteristics: Trimmed Samples

| | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| | Is a Manager in 5 Years | | | Promoted within 5 Years | | |
| | Overall | Lower Bound | Upper Bound | Overall | Lower Bound | Upper Bound |
| *Panel A: All Mothers* | | | | | | |
| Eligibility*Infant | -0.0943 | -0.0893 | -0.0954 | -0.0657** | -0.0602* | -0.0704** |
| | (0.0618) | (0.0636) | (0.0636) | (0.0305) | (0.0309) | (0.0307) |
| Job Protection*Infant | -0.0646* | -0.0506 | -0.0691* | 0.0213 | 0.00158 | 0.0358 |
| | (0.0386) | (0.0375) | (0.0393) | (0.0399) | (0.0419) | (0.0414) |
| Observations | 3,625 | 3,544 | 3,544 | 3,454 | 3,373 | 3,373 |
| R-squared | 0.798 | 0.785 | 0.787 | 0.648 | 0.704 | 0.704 |
| *Panel B: Mothers with High Education* | | | | | | |
| Eligibility*Infant | -0.0518 | -0.0478 | -0.0504 | 0.00267 | -0.00204 | 0.0103 |
| | (0.223) | (0.224) | (0.284) | (0.0805) | (0.0835) | (0.0839) |
| Job Protection*Infant | -0.125*** | -0.0979* | -0.134*** | -0.0749* | -0.0690* | -0.0986*** |
| | (0.0521) | (0.0511) | (0.0524) | (0.0402) | (0.0399) | (0.0397) |
| Observations | 1,634 | 1,620 | 1,620 | 1,576 | 1,563 | 1,563 |
| R-squared | 0.797 | 0.796 | 0.798 | 0.710 | 0.712 | 0.708 |
| *Panel C: Mothers with Low Education* | | | | | | |
| Eligibility*Infant | -0.0509 | -0.0489 | -0.0733 | -0.106*** | -0.0898* | -0.129** |
| | (0.152) | (0.152) | (0.151) | (0.0451) | (0.0499) | (0.0520) |
| Job Protection*Infant | 0.0637 | 0.0491 | 0.0864 | 0.0633 | 0.0476 | 0.0880 |
| | (0.111) | (0.114) | (0.105) | (0.0917) | (0.0962) | (0.0905) |
| Observations | 1,577 | 1,541 | 1,541 | 1,502 | 1,466 | 1,466 |
| R-squared | 0.747 | 0.748 | 0.750 | 0.704 | 0.706 | 0.709 |

Cluster-robust standard errors in parentheses. Each column in each panel is a separate regression comparing mothers of infants to mothers whose youngest child is 3-4 years old in the reference year. Controls include age, marital status, number of kids, indicators for age of youngest child, and race and education dummy variables. All regressions time fixed effects and individual fixed effects as well. To calculate the lower bound of the effect on each outcome, I drop $\beta$ percent of the observations in the group of mothers of infants in the relevant policy period that all have a value of 1 for the outcome variable. The upper bound effect drops $\beta$ percent of the observations that all have a value of 0 for the outcome. $\beta$ is the coefficient from the regression of the effect of each policy on employment in five years, and varies by education.

*** p<0.01, ** p<0.05, * p<0.1