THE ECONOMIC CONSEQUENCES OF PARENTAL LEAVE MANDATES: LESSONS FROM EUROPE*

CHRISTOPHER J. RUHM

This study investigates the economic consequences of rights to paid parental leave in nine European countries over the 1969 through 1993 period. Since women use virtually all parental leave in most nations, men constitute a reasonable comparison group, and most of the analysis examines how changes in paid leave affect the gap between female and male labor market outcomes. The employment-to-populations ratios of women in their prime childbearing years are also compared with those of corresponding aged men and older females. Parental leave is associated with increases in women's employment, but with reductions in their relative wages at extended durations.

Over 100 countries have enacted some form of parental leave policies, with most assuring at least two to three months of paid job absences [Kamerman 1991]. Nevertheless, the effects of providing rights to time off work in the period surrounding childbirth remain poorly understood. Proponents believe that parental leave results in healthier children and improves the position of women in the workplace. Opponents counter that the mandates, by restricting voluntary exchange between workers and employers, reduce economic efficiency and may have a particularly adverse effect on women.

The results of previous research on parental leave are ambiguous. Some U. S. studies suggest that time off work is associated with increases in employment and wages [Dalto 1989; Spalter-Roth and Hartmann 1990; Waldfogel 1994, 1997]. However, since these analyses cover a period when most leaves were voluntarily provided by employers, rather than being required by law, the differences in labor market status may result from nonrandom selection into jobs providing the benefit, and the evidence is difficult to interpret. Other researchers have attempted to overcome the selection problem by examining legislated parental leave benefits. Klerman and Leibowitz [1997] uncover mixed

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employment effects of maternity leave mandates instituted by some states during the late 1980s. Waldfogel [1996] finds that recently enacted federal legislation in the United States had little effect on wages, while modestly increasing employment; but this last result is sensitive to the model estimated. The ambiguous results of these studies may reflect the limited scope of the federal and state mandates or inadequacies of the data. Finally, Ruhm and Teague [1997], using information for seventeen nations, show that short to moderate entitlements to parental leave are positively related to per capita incomes, employment-to-population ratios (EP ratios), and labor force participation rates. However, there is little indication of stronger effects for women than for men, raising concern that the direction of causation may be misidentified.

This study investigates the labor market consequences of rights to paid parental leave using data for nine European countries over the 1969 through 1993 period.[1] The dependent variables are EP ratios and hourly wages.[2] Since women use virtually all parental leave in most countries, men constitute a reasonable comparison group, and the "natural" experiment examines how changes in leave entitlements affect the gap between female and male outcomes.[3] Limited analysis is also undertaken using 25–34 year old women as the treatment group and corresponding men or females aged 45–54 as the comparison group. The younger women are in their prime childbearing years and so should be strongly affected by leave mandates. Time and country effects are controlled for throughout the analysis to provide "difference-in-difference-in-difference" (DDD) estimates. Country-specific time trends are frequently included to capture the effects of group-specific factors that vary over time within countries.

European data are particularly useful for investigating the effects of parental leave. All Western European countries cur-

1. A distinction is sometimes made between "maternity leave," which is granted to mothers for a limited period around the time of childbirth, and "parental leave," which permits additional time off work to care for infants or young children. Both are included in the definition of parental leave used below.

2. An earlier version of this paper also included weekly work hours as an outcome. There was little indication of a strong parental leave effect, and the results were sensitive to the specification chosen, probably partly because sex-specific data on work hours were unavailable for many countries.

3. Gruber [1994] and Waldfogel [1996] have similarly used men as a comparison group when examining the effects of mandated maternity benefits and parental leave legislation in the United States.

rently offer at least three months of paid maternity benefits, but many of the policies have been instituted or significantly revised during the last 30 years, resulting in substantial variation over time and across countries in the type and duration of the entitlements. Conversely, the United States did not require employers to provide parental leave until the 1993 passage of the Family and Medical Leave Act (FMLA).[4]

Better understanding the effects of parental leave mandates is important in both the European and United States contexts. Europe has been grappling with the question of whether extensive social protections inhibit economic flexibility and are a cause of low rates of recent employment growth [Blank 1994]. These concerns have recently led a number of countries to shorten the period of leave or reduce payments provided during it, at the same time that other nations have increased them [Organization for Economic Cooperation and Development 1995]. Conversely, advocates (e.g., the Carnegie Task Force on Meeting the Needs of Young Children [1994]) have argued for broadening the U. S. federal law to include small employers and provide payment during the time off work.

To preview the results, rights to paid leave are found to raise the percentage of women employed, with a substantial effect observed for even short durations of guaranteed work absence. In the preferred econometric specifications, leave legislation raises the female employment-to-population ratio by between 3 and 4 percent, with larger effects for women of childbearing age. Around one-quarter of this change probably results from increases in the number of women who are reclassified as "employed but absent from work" due to the availability of leave. Brief leave entitlements have little effect on women's earnings, but lengthier leave is associated with substantial (2 to 3 percent) reductions in relative wages.

4. The FMLA requires employers with more than 50 workers in a 75-mile area to allow twelve weeks of unpaid leave following the birth or adoption of a child or for personal illness or the health problem of a family member. Health insurance contributions must be continued during the period. Firms need not provide leave to the highest paid 10 percent of their workforce or persons employed less than 1250 hours during the previous year [Ruhm 1997]. Ten states and the District of Columbia legislated job-protected work absences prior to the FMLA, and eight others supplied limited rights to parental leave without guaranteeing the reinstatement of employment [Waldfogel 1994]. The state laws were enacted in the late 1980s or early 1990s and included numerous exemptions.

I. The Economics of Parental Leave Mandates

In a competitive spot labor market with perfect information and no externalities, mandated benefits such as parental leave reduce economic efficiency by limiting the ability of employers and workers to voluntarily select the optimal compensation package. Nevertheless, supporters argue that parental leave entitlements improve the health and well-being of children (e.g., Zigler, Frank, and Emmel [1988] and the Carnegie Task Force on Meeting the Needs of Younger Children [1994]). This might occur if the benefits represent externalities that are not adequately valued by agents negotiating labor contracts. For instance, the gains might not be fully taken into account if workers have inadequate information concerning the advantages of staying at home with infants, if they pay only a portion of the costs of their children's medical care (as with most types of health insurance), or if they have higher than socially optimal discount rates. Employers may also be less aware or supportive of the advantages of parental leave to dependents than of the corresponding benefits to the workers themselves.

It is also frequently asserted that leave mandates decrease female unemployment and increase firm-specific human capital by reducing the need for women to change jobs, if they wish to spend time at home with young children [Kamerman 1988; Bookman 1991; Bravo 1991; Trycinski 1991]. Lacking some source of market failure, this argument is unconvincing. Employers and workers can always voluntarily negotiate maternity leave, mitigating the joblessness and retaining the specific investments. Moreover, with competitive labor markets, the groups most likely to use parental leave will pay for it by receiving lower wages, implying that females of childbearing age will continue to obtain lower and possibly reduced compensation if the benefit is mandated.[5] Entitlements that allow substantial time off work may cause employers to limit women to jobs where absences are least costly, thereby increasing occupational segregation, as Stoiber [1990] suggests has occurred in Sweden.

Adverse selection under asymmetric information provides a potential source of market failure. A company voluntarily offering leave is likely to attract a disproportionate number of "high-risk" employees and be forced to pay lower wages. Persons with small

5. See Gruber [1994] for an excellent discussion of group-specific mandates and Summers [1989], Mitchell [1990], or Krueger [1994] for more general discussions of the economics of mandated benefits.

probabilities of using the benefit will avoid these firms and so do without even socially optimal leave.[6] A government mandate eliminates the incentive for this type of sorting behavior and has the potential to raise welfare.[7]

Companies in the United States rarely provided explicit paid maternity leave prior to the FMLA. Only 3 percent of full-time employees in private medium and large establishments (greater than 100 workers) were entitled to such leaves in 1993 and 1 percent of those working for small employers in 1992 [U. S. Department of Labor, Bureau of Labor Statistics 1994a, 1994b]. These low coverage rates could indicate that the costs of the entitlements exceed the benefits or that market imperfections limit their unregulated provision. Alternatively, most workers may have been able to take time off work through vacation, sick leave, or temporary disability policies, even without explicit maternity leave.[8]

Parental leave mandates are likely to shift the labor supply curve of the groups most probable to use it to the right (relative to those workers less likely to take leave).[9] The demand curve simultaneously moves to the left. However, since leave benefits are paid primarily by the government in most European countries, demand only shifts to the extent that nonwage costs (e.g., expenses associated with hiring and training temporary replace-

6. This is analogous to Rothschild and Stiglitz's [1976] argument for market failure in insurance markets. Aghion and Hermalin [1990] suggest that in some situations socially optimal parental leave might not be voluntarily provided to any workers. In their model, low-risk individuals signal this to employers by agreeing to contracts providing for little or no leave. High-risk workers sometimes do better by mimicking their counterparts, by taking positions without leave, than by revealing their propensity toward absenteeism.

7. The inefficiency of privately negotiated labor contracts under asymmetric information has been demonstrated across a variety of contexts. For example, McGuire and Ruhm [1993] indicate that employer-drug testing is likely to be excessive, and Levine [1991] and Kuhn [1992] argue that just-cause employment security regulations and advance notice of job terminations may be under-provided.

8. The Pregnancy Discrimination Act (PDA) of 1978 requires companies offering leave for temporary disabilities, which includes most medium and large establishments, to cover pregnancy and childbirth in the same way as other temporary disabilities. Several states have supplemented the PDA with stronger temporary disability laws or maternity leave mandates [Ruhm 1997]. During the 1986–1988 period, 73 percent of "employed" women in the United States with one-month old infants were on leave (and 41 percent on paid leave) rather than working, as were 41 percent (16 percent) of those with two-month old babies [Klerman and Leibowitz 1994].

9. In particular, some individuals will increase their labor supply prior to having children in order to meet the qualification conditions for parental leave. I return to this point below. Mortensen [1977] makes an analogous argument with regard to unemployment insurance.

ments) increase.[10] Thus, the shift in supply is likely to be large compared with that in demand, implying that the relative employment of women will rise and their relative wages will fall in the new equilibrium.[11]

Increased leave-taking could reduce work in the period immediately surrounding childbirth, even if leave entitlements raise overall employment. However, Klerman and Leibowitz [1997] illustrate that employment may increase even during this time span. The reason is that some persons who would otherwise have terminated their jobs to take more leave than previously permitted, may now find it worthwhile to return to work sooner in order to remain with their old employers. This occurs because the gap between desired leave duration and that offered by the firm decreases, while the benefits of maintaining the employment relationship (e.g., higher future compensation) are little changed.

There could be additional "dynamic" effects. For instance, labor productivity will rise if parental leave increases firm-specific human capital by allowing individuals to return to their old jobs. This will shift the demand curve to the right, further increasing employment and attenuating or reversing the decline in wages. Alternatively, if human capital depreciates during lengthy leave periods, the employment increases will be smaller, and the earnings reductions larger than in the static case.

II. Parental Leave Policies in Europe

Legislated maternity benefits have a long history in Europe. The German Imperial Industrial Code of 1891 set maximum work hours and *prohibited* the employment of women within four weeks of childbirth. Amendments to the code in 1903 and 1911 increased the leave period to six weeks and supplied women with paid time off work in the two weeks before delivery. By the turn of the century there was discussion of providing maternity insurance in many European countries.[12] Most early legislation emphasized concern for the health of the child and mother. Prenatal and postnatal leave was typically compulsory, and income support or

10. More precisely, this refers to the movement of the demand curve compared with groups not using leave. The demand for all types of labor may decline if the parental leave benefits are financed by payroll taxes levied on employers.

11. If the group-specific mandate is imposed in the presence of binding equal pay legislation or union rules that restrict wage reductions, female employment is likely to rise less (or may even fall), the decline in wages will be smaller, and the deadweight loss is likely to be larger [Gruber 1994].

12. See Frank and Lipner [1988] or Teague [1993] for discussions of early maternity leave policies.

job-protection was seldom provided. The 1919 and 1952 International Labour Organization *Maternity Protection Conventions* recommended that women not be permitted to work during the six-week period following confinement. Payment during leave and rights to return to the old job were also advocated, but many countries did not adopt these suggestions until much later.[13]

After the end of World War II, many nations that had recruited women into previously male-dominated occupations wished to return them to the home [Moeller 1993]. The motivation for policies related to family allowances, protective legislation, and family-law reform was often to restore women to their "proper" roles as mother and wife [Frank and Lipner 1988]. In the postwar period some countries mandated compulsory pregnancy leave but failed to prohibit dismissal from jobs.

By the late 1960s the concept of maternity leave began to evolve from a prohibition on employing women during the period surrounding childbirth to one of job-protected time off work to care for newborns and young children. Portugal, Spain, and Finland instituted employment reinstatement provisions during the 1969–1971 period; France and the Netherlands passed similar legislation in 1975 and 1976; as did Denmark, Ireland, and Greece between 1980 and 1984. Other nations, such as Switzerland and the United Kingdom, inaugurated regulations providing job-protected maternity leave. Nonetheless, vestiges of protective legislation still persist in some countries. Postnatal leave remains compulsory, rather than voluntary, in many nations, and some (such as Austria, France, and Italy) continue to require prenatal leave [Brocas et al. 1990].

Income support is now provided during at least a portion of the work absence throughout Europe. Wage replacement rates often exceed 80 percent and are typically financed by a combination of payroll taxes and general government revenues, although some nations require direct employer contributions. Although a few countries have recently reduced replacement rates or leave durations, the overall trend has been toward longer leave periods, with fathers increasingly gaining rights to time off work [Organization of Economic Cooperation and Development 1995]. The European Community Social Charter recently established a mini-

13. The 1919 convention advocated twelve weeks of paid leave and job-reinstatement upon return to employment. The 1952 conference recommended a cash benefit equal to at least two-thirds of previous earnings, compared a previously suggested unspecified amount "sufficient for the full and healthy maintenance of the working mother and her child" [International Labor Office 1984].

mum standard leave period of fourteen weeks, with pay no less than the individual would receive if absent from work because of sickness [Addison and Siebert 1993].

Even where parental leave extends to fathers, mothers take the vast majority of time off work.[14] There are a variety of reasons why men take leave so sparingly. In addition to cultural norms and differences in earnings capacity, the entitlements are generally restricted to mothers during the period immediately surrounding confinement, and fathers often can subsequently take time off work only if the mother qualifies for but waives her rights to it.

III. Data

This analysis uses aggregate data covering the 1969 through 1993 period for nine European countries (Denmark, Finland, France, Germany, Greece, Ireland, Italy, Norway, and Sweden). Labor market data for years prior to 1969 are frequently incomplete, and parental leave policies changed little during the early and middle 1960s. The nine nations chosen are all Western European countries with significant changes in their paid parental leave policies during the sample period.[15]

Paid parental leave is defined to include rights to time off work during the period surrounding childbirth where the size of the income support is directly related to previous employment. This is distinguished from payments that are available to all individuals, regardless of their work histories. Most of the analysis focuses on job-protected leave, where dismissal is prohibited during pregnancy and job-reinstatement is guaranteed at the end of the leave, since employment security is likely to be a key characteristic of leave policies that workers consider a "benefit." During the period analyzed, several nations added job security provisions to previously enacted compulsory "maternity protection" laws.

14. Even in Sweden, which provides the strongest encouragement for men to take some leave, males accounted for just 7 percent of total weeks of parental leave in 1988 [Organization of Economic Cooperation and Development 1995]. More typical is Germany, where fewer than 1 percent of those receiving parental leave in 1989 were men [Der Bundesminister fuer Jugend, Familie, Frauen und Gesundheit 1989]. (I thank Katharina Spiess for providing me with and translating this information.)

15. Eight other countries (Austria, Belgium, Canada, the Netherlands, Portugal, Spain, Switzerland, and the United Kingdom) were included in an earlier investigation of paid and unpaid leave by Ruhm and Teague [1997]. Canada was excluded from this analysis in order to restrict the sample to European nations. The other seven nations were deleted because they had little or no change in paid leave entitlements during the sample period.

A measure of "full-pay" weeks of leave is calculated by multiplying the number of weeks of paid leave by the average wage replacement rate during the period. The replacement rates are approximations because they do not account for minimum or maximum payments which sometimes exist. Also, some nations provide a "flat rate" payment or a fixed payment plus a percentage of earnings. In these cases, the replacement rate was estimated as a function of average female wages.

The leave durations apply to persons meeting all eligibility criteria. This overstates actual time off work, since some individuals do not fulfill the employment requirements and others use less than the allowed absence. Qualifying conditions either have not changed or have loosened over time in most countries, and increased labor force participation rates imply that more women are likely to meet given work requirements. Therefore, a greater share of females are expected to qualify for benefits at the end of the period than at the beginning, implying that the secular increase in parental leave entitlements is understated.

Unpaid leave has not been incorporated into this analysis for two reasons. First, many employers may be willing to grant unpaid time off work, even in the absence of legislation, making it difficult to distinguish between the effects of job absences voluntarily granted by employers and those required by law. Second, the actual use of legislated rights to unpaid leave may be quite limited, particularly for the extremely lengthy entitlements now provided in some countries. I also do not distinguish between leave available only to the mother and that which can be taken by either parent. Nor do I model differences in "take-up" rates.

These restrictions should be kept in mind when interpreting the results. If (within-country) growth in paid leave entitlements is positively correlated with changes in either the proportion of women with qualifying work histories or rights to additional unpaid leave, the econometric estimates will represent the combined effects of these factors, and so will overstate the impact of an increase in paid leave which occurs in isolation.

Data for Germany are only included through 1985. In 1986, Germany simultaneously lengthened the duration of job-protected leave and extended to nonworkers the income support payments previously restricted to persons meeting qualifying employment conditions [Ondrich, Spiess, and Yany 1996]. Using the criteria discussed above, this would be defined as a reduction in paid leave, since the payments are no longer tied to previous employment. Such a classification seems problematic, given that the duration of job-protected time off work was substantially increased in 1986 and again in 1988. The easiest way of dealing with the problem was to delete observations after 1985.[16]

Information on parental leave was obtained from the International Labour Office's *Legislative Series*, their 1984 global survey on "Protection of Working Mothers," and from *Social Security Programs Throughout the World*, which is published biennially by the United States Social Security Administration. A subset of the leave data was previously used by Teague [1993] and Ruhm and Teague [1997]. The time period has been extended in the present paper, and the data have been rechecked and modified as appropriate for greater accuracy.

The dependent variables are (natural logs of) employment-to-population ratios and hourly wage rates. Data on EP ratios are from various issues of the OECD publication *Labour Force Statistics*; those on wages are from several volumes of the ILO *Yearbook of Labour Statistics*.[17] Employment information is available for all nine countries but gender-specific wage data are more difficult to obtain. This analysis uses information on the wages of manufacturing workers for six nations (Finland, France, Greece, Ireland, Norway, and Sweden), for all nonagricultural workers in the case of Denmark and Germany, and with no wage data obtained for Italy. The frequent restriction to manufacturing implies that the results for wages should be interpreted cautiously. Nominal wages were deflated by purchasing power parities, using OECD *National Accounts* data, and by the Consumer Price Index. Age-specific information on the EP ratios of 25–34 and 45–54 year old women and 25–34 year old men was also used for all countries except Greece and Norway.[18] Finally, demographic data were obtained from *Labour Force Statistics* on population (of civilians aged 15–64), birth rates (per 1000 resident

16. As an alternative, I estimated a preliminary set of models with observations for Germany included through 1990 (the year before German Unification) and parental leave entitlements assumed to either remain constant (at 32 weeks) after 1985 or to increase in accordance with the extensions granted in 1986 and 1988. In the first case, the estimated parental leave effects were virtually identical to those obtained when the post-1985 data were excluded. In the second, the predicted increases (decreases) in female EP ratios (wages) were slightly larger (smaller).

17. The employment-to-population ratio is calculated as civilian adult employment divided by the population between the ages of 15 and 64, using standardized OECD definitions. Wages generally refer to hourly straight-time pay (either wages or salaries), excluding overtime premiums, bonuses, or gratuities, and averaged over both full- and part-time workers.

18. The data for Italy refer to 25–39 and 40–49 year olds.

population), total unemployment rates, and the proportion of the working age population employed in service or agricultural jobs (with manufacturing the excluded reference category).

The data are not always completely comparable between or within countries. For example, purchasing power parities provide the best method of adjusting nominal wages but are unlikely to supply exactly equivalent information across time and place. Similarly, nations sometimes alter their methods of collecting or aggregating data. The estimation strategy is designed to minimize biases resulting from such noncomparabilities. The inclusion of country fixed-effects controls for differences (between countries) in collection methods that remain constant over time. Most of the analysis focuses on gender differences in labor market outcomes. This automatically accounts for breaks in series (within-countries) which have the same effect on the male and female aggregates. Examination of the ratio of (the log of) female-to-male outcomes for periods immediately preceding and following each interruption in series revealed only one case, Norway in 1971, where the break led to a substantial change in the relative size of male and female EP ratios. Norwegian data for the years 1969 and 1970 were therefore deleted from the analysis.

IV. Time-Trends

Parental leave entitlements increased sharply between 1969 and 1993. Weighting observations by the country's working age population, the mean duration of paid leave for the eight countries (excluding Germany for which no data were collected after 1985) rose from 10 to 33 weeks while average full-pay weeks grew from 7 to 22 weeks (see Fig Ia). The increases were most dramatic during the first ten years of the period, with a particularly large jump occurring at the end of the 1970s when six countries (Finland, France, Germany, Italy, Norway, and Sweden) almost simultaneously raised entitlements to job-protected leave. Since 1980, there has been little overall rise in leave durations, as increases in some countries have offset declines in others. Full-pay weeks grew more slowly than partially paid leave because some of the additional entitlements to time off work were provided at relatively low wage replacement rates.

Table I summarizes parental leave provisions in the last year of the data (1993 except for Germany). At that time, the countries offered a minimum of fourteen weeks of paid leave, and

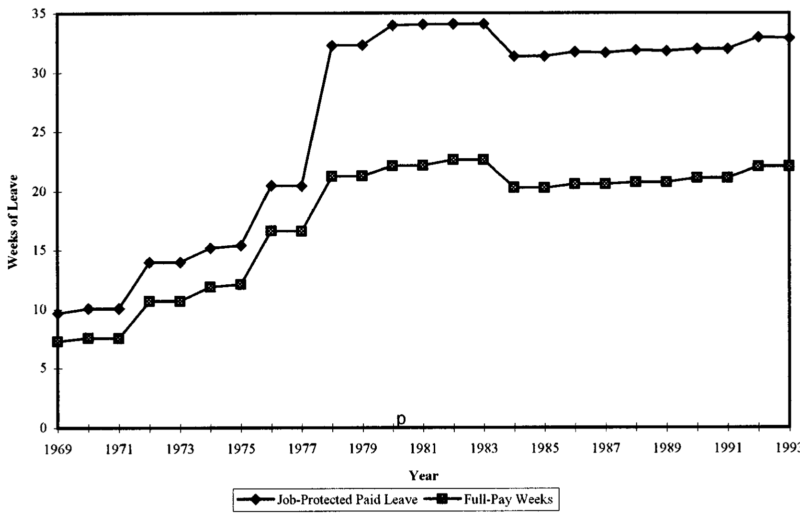

FIGURE Ia
Average Weeks of Parental Leave

six nations provided rights to more than six months off work. Full-pay weeks ranged from 9 weeks in Greece to 58 weeks in Sweden, with a positive correlation between replacement rates and leave durations. Income support during the work absence was typically financed through a combination of payroll taxes and general revenues. The conditions required to qualify for leave varied, but persons with more than a year of service were usually covered.

Table II displays paid leave durations and estimated replacement rates for each country at four-year intervals. The number of nations providing some job-protected paid entitlement rose from four in 1969 to eight in 1977, with all nine doing so after 1983. Countries supplying parental benefits in 1969 extended them during the sample period, with the result that the dispersion of leave durations tended to increase over time.[19] There were 30 observed changes in durations over the sample period and 5

19. The standard deviation of the (population-weighted) duration of paid leave (full-pay weeks) for the eight countries other than Germany was 10.8, 22.7, and 18.0 (8.6, 11.5, and 12.4) weeks in 1969, 1981, and 1993, respectively. The difference between minimum and maximum entitlement was 21, 57, and 50 (17, 37, and 49) weeks in these same years.

TABLE I
PAID PARENTAL LEAVE IN 1993

Country	Amount of leave	Rate of pay	Source of funds	Qualification conditions
Denmark	28 weeks	90% with maximum	Employers, Government	120 hours of employment during preceding 3 months.
Finland	44 weeks	80% with minimum; lower rate at high incomes	Payroll taxes, Government	Residence in country.
France	16 weeks	84% with minimum and maximum	Payroll and dedicated taxes	Insured 10 months before leave; minimum work hours or insurance contributions.
Germany	32 weeks	100% with minimum and maximum	Payroll taxes, Government	12 weeks of insurance or 6 months of employment.
Greece	15 weeks	60% with minimum	Payroll taxes, Government	200 days of contributions during last 2 years.
Ireland	14 weeks	70% with maximum	Payroll taxes, Government	39 weeks of contributions.
Italy	48 weeks	53% (80% first 5 months; 30% next 6 months)	Payroll taxes, Government	Employed and insured at start of pregnancy.
Norway	42 weeks	100% with maximum	Payroll taxes, Government	Employed and insured at least 6 of the last 10 months.
Sweden	64 weeks	90%	Payroll taxes, Government	Insured 240 days before confinement.

Information for Germany refers to 1985.

additional cases where nations modified replacement rates without altering the length of leave.

The relative wages and EP ratios of women also rose over time. This is shown in Fig Ib, which displays the (population-

TABLE II
ENTITLTEMENTS TO JOB-PROTECTED PAID PARENTAL LEAVE (IN WEEKS) AND WAGE REPLACEMENT RATES IN SELECTED YEARS

	1969	1973	1977	1981	1985	1989	1993
Denmark	0	0	0	18	28	28	28
				[.90]	[.90]	[.90]	[.90]
Finland	0	12	29	43	43	44	44
		[.55]	[.55]	[.55]	[.80]	[.80]	[.80]
France	0	0	14	16	16	16	16
			[.90]	[.90]	[.90]	[.90]	[.84]
Germany	14	14	14	32	32		
	[1.00]	[1.00]	[1.00]	[1.00]	[1.00]		
Greece	0	0	0	0	12	12	15
					[.60]	[.60]	[.60]
Ireland	0	0	0	12	14	14	14
				[.55]	[.70]	[.70]	[.70]
Italy	21	21	31	57	48	48	48
	[.80]	[.80]	[.80]	[.57]	[.53]	[.53]	[.53]
Norway	12	12	12	18	18	24	42
	[.13]	[.32]	[.32]	[1.00]	[1.00]	[1.00]	[1.00]
Sweden	16	16	30	52	52	52	64
	[.55]	[.55]	[.90]	[.71]	[.71]	[.90]	[.90]

The unbracketed entry shows the number of weeks of job-protected paid parental leave employers are required to offer. The bracketed entry displays wage replacement rates. These are sometimes subject to minimum or maximum amounts. The average replacement rate is estimated in cases where the income support during early and later portions of the leave or includes a flat rate payment.

weighted) female-to-male ratios of these outcomes for the eight countries other than Germany.[20] Women were employed less than half as often as men (40 percent versus 82 percent) in 1969 but worked 70 percent as frequently (47 percent versus 67 percent) in 1993; the earnings gap fell from 25 percent to under 20 percent during the same period. Since many factors other than parental leave will have contributed to the declining gender differentials, it is important that the econometric methods control for these sources of spurious correlation.

One issue deserving mention is that European countries often count individuals on parental leave as "employed but absent from work" rather than "not employed." The extension of leave entitlements will therefore raise reported EP ratios if the work absences of "employed" persons increase. As discussed below, this is likely to

20. Specifically, the figure shows $(EP_f/EP_m) - 1$ and $(W_f/W_m) - 1$, for EP_i and W_i representing the EP ratios and wages of the ith group, with i equal to f for females and m for males.

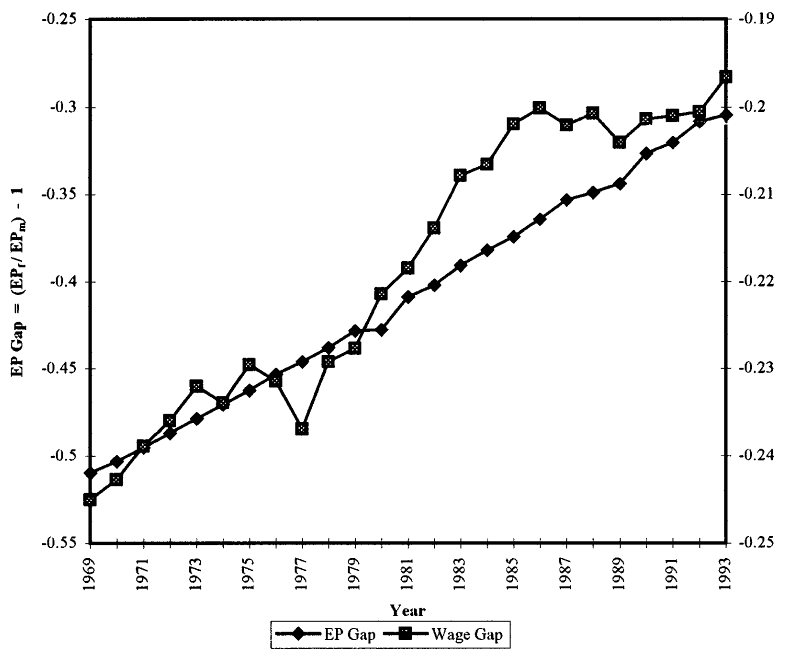

FIGURE Ib
Gender Gap in EP Ratios and Wages

account for one-quarter to one-half of the observed positive relationship between parental leave durations and EP ratios.

V. ESTIMATION STRATEGY

Labor market outcome Y, measured in natural logs, for sex group i (where f indicates females and m males) in country j at year t, is assumed to be determined by

$$(1)\quad Y_{ijt} = a_1S_i + a_2C_j + a_3T_t + b_1(S_i \times C_j) + b_2(S_i \times T_t) + b_3(C_j \times T_t) + d_iL_{jt} + e_{ijt}.$$

The key dependent variable L_{jt} is weeks of paid parental leave entitlement. S_i is a group-specific intercept, C_j a country effect, T_t a time effect, and e_{ijt} an i.i.d. error term. The second level interactions allow for sex-specific country and time effects and for a general time-varying country effect. Other observables are excluded for ease of exposition.

The sex difference in outcomes can be expressed as

$$(2)\quad Y_{fjt} - Y_{mjt} = a_1(S_f - S_m) + b_1(S_f - S_m)C_j + b_2(S_f - S_m)T_t + (d_f - d_m)L_{jt} + (e_{fjt} - e_{mjt})$$

or equivalently

$$(3)\quad \Delta Y_{jt} = \alpha + \beta_1 C_j + \beta_2 T_t + \delta L_{jt} + \epsilon_{jt}$$

Equation (3) is a DDD model. β_1 and β_2 indicate gender-specific country and time differences; δ shows the sex-difference in the impact of parental leave. Thus, these estimates measure how growth in the gender gap in labor market outcomes varies as a function of within-country changes in leave entitlements.

Since women use almost all parental leave, it may be reasonable to assume that $d_m = 0$. In this case δ supplies an unbiased estimate of d_f. By contrast, if d_m is nonzero and has the same (opposite) sign as d_f, the regression coefficient will be biased toward (away from) zero. One reason why d_m and d_f might have the same sign is that some men do use some parental leave. However, the resulting bias is likely to be small, since males take only a tiny fraction of total weeks of leave in most countries. Conversely, d_m and d_f will have the opposite sign if employers or households respond to lengthened entitlements by substituting employment away from females and toward males (or away from younger and toward older women) or vice versa. In this case, δ provides an upper bound estimate of d_f.

It is useful to contrast the DDD model to the corresponding equation without a comparison group:

$$(4)\quad Y_{ijt} = \alpha + \beta_1 C_j + \beta_2 T_t + \delta L_{jt} + \epsilon_{ijt},$$

where $\alpha = a_1 S_i$, $\beta_1 = a_2 C_j + b_1 (S_i \times C_j)$, $\beta_2 = a_3 T_t + b_2 (S_i \times T_t)$, and $\epsilon_{ijt} = b_3 (C_j \times T_t) + e_{ijt}$. Equation (4) examines within-country growth in the dependent variable as a function of modifications in leave durations, but it does not contrast these changes to those of a comparison group expected to be unaffected by the leave entitlements (the third difference in the DDD model). Bias will therefore be introduced if time-varying country-specific effects $(C_j \times T_t)$ are correlated with changes in parental leave, as might occur if nations choose to increase entitlements when employment is rising.[21]

21. If $d_m = 0$, the parental leave coefficient in equation (4), estimated for male outcomes, provides a direct indication of this bias.

The DDD specification in equation (3) accounts for time-varying factors that affect both sexes equally. However, the estimates may still be inconsistent if within-country changes in parental leave are correlated with unobservables that have different effects on female and male labor market outcomes. This can be seen by adding a third level interaction $c_1(S_i \times C_j \times T_t)$ to equation (1). The error term in (3) becomes $\epsilon_{ijt} = c_1 (S_f - S_m)(C_j \times T_t) + (e_{fjt} - e_{mjt})$, which may be correlated with L_{jt}. Omitted explanatory variables represent a potentially important source of sex-specific time-varying factors.[22] A vector of country-specific time trends will therefore frequently be added to the models. These eliminate the bias, to the extent that sex differentials in omitted characteristics follow the specified trend.

To adjust for heteroskedasticity resulting from differences in population sizes, most of the models are estimated by weighted least squares (WLS). The weights are determined using the following procedure. First, the equations are estimated by OLS. Second, the squared residuals from these models are regressed against a constant term and the reciprocal of the working age population. Finally, the square root of the inverse of the predicted values from the second-stage regression are used as weights in the final set of estimates. Blackburn [1995] shows that this procedure is more efficient than weighting by (the square root of) population size or using OLS and reporting Huber-White standard errors, if there is a common group effect or group-time interaction across individuals in a country.[23]

VI. Results

A first set of econometric estimates is displayed in Table III. Vectors of country and time dummy variables are included, observations are weighted to adjust for heteroskedasticity using the procedure described above, and the leave regressor is weeks of job-protected paid entitlement (irrespective of the replacement rate) divided by 100. The first two rows of each panel show results of equation (4), separately estimated for males and females. The dependent variable in the third row is the difference between

22. For instance, if the education of women is rising relative to men in countries extending leave, increased schooling could induce a spurious positive correlation between leave durations and female employment or earnings.

23. A large and significant constant term is obtained in virtually all of the second-stage regressions, which confirms that such group effects are important and justifies the use of this weighting procedure.

TABLE III
ECONOMETRIC ESTIMATES OF PARENTAL LEAVE EFFECTS USING LINEAR SPECIFICATIONS

Group	(a)	(b)	(c)
	Employment-to-population ratio ($n = 203$)		
Females (Y_{fjt})	.2550	.1372	.0564
	(.0650)	(.0528)	(.0485)
Males (Y_{mjt})	.1887	.0494	−.0028
	(.0449)	(.0299)	(.0386)
Difference ($Y_{fjt} - Y_{mjt}$)	.0583	.0859	.0597
	(.0467)	(.0426)	(.0320)
	Hourly wages ($n = 173$)		
Females (Y_{fjt})	−.3923	−.1993	−.0757
	(.0977)	(.0885)	(.0791)
Males (Y_{mjt})	−.3197	−.2159	−.0119
	(.0823)	(.0813)	(.0749)
Difference ($Y_{fjt} - Y_{mjt}$)	.0152	.0426	−.0624
	(.0400)	(.0380)	(.0303)
Demographics	No	Yes	No
Time trends	No	No	Yes

The table displays coefficients on parental leave regressors. Data are for nine European countries over the 1969–1993 period. All specifications include vectors of year and country dummy variables. Parental leave refers to weeks of job-protected paid entitlement (irrespective of the wage replacement rate) divided by 100. Demographic variables include birthrates, unemployment rates, and the fraction of employment in service and agricultural jobs. Dependent variables are natural logs of female and male labor market outcomes in the first two rows of each panel and differences in the natural log of female and male outcomes in the third. Observations are weighted to correct for heteroskedasticity. Standard errors are in parentheses.

female and male outcomes; hence this is the DDD model specified by equation (3).

Demographic characteristics (birthrates, unemployment rates, and the employment shares in agriculture and services) are controlled for in column (b), as are country time trends in column (c). The demographic variables are likely to capture a portion of the impact of time-varying factors that influence the gender gap in EP ratios or wages. However, some of these regressors may be endogenous (e.g., countries may use parental leave policies as part of a strategy to raise birthrates). Moreover, since the set of characteristics controlled for is quite limited, the inclusion of country time trends may more adequately proxy the sex-specific time-varying effects.

The table demonstrates the importance of including a comparison group. The leave coefficients are of roughly similar magnitude for men and women in specification (a), with the result that the DDD coefficient in the third row is small and insignificant for both EP ratios and wages. Demographic variables capture the effects of some of the confounding factors, as evidenced by the reduction in the absolute value of the leave coefficients in the first two rows of each panel in column (b). However, the parameter estimates decrease substantially more in column (c), suggesting that the country-specific time trends do a better job of accounting for sex-specific time-varying effects. Indeed, the leave coefficients for the male outcomes are small and statistically insignificant in this specification, which is consistent with the hypothesis that parental leave has no effect on the male labor market. The DDD estimates in the third row are fairly insensitive to the choice of regressors for EP ratios but show more variation for wages.

A. Quadratic Specifications

Table IV summarizes the results of DDD models where the dependent variable is the difference between (the log of) female and male outcomes, LEAVE indicates weeks of job-protected paid leave divided by 100, and a quadratic term is included to allow for nonlinearities. The *p*-value refers to the null hypothesis that the parental leave coefficients (LEAVE and its square in these regressions) are jointly equal to zero. The lower panel shows the predicted impact of specified paid leave entitlements, compared with the case of no mandate.[24] Models (c) and (d) differ in that the latter is estimated by OLS, but with Huber-White robust standard errors reported, rather than using the WLS procedure.

Paid leave is positively related to the percentage of females employed, but there is some evidence that lengthy entitlements reduce their relative wages. The null hypothesis of no parental leave effect is not rejected in model (a), but the estimates are large and significant when time trends are controlled for (specification (c)); intermediate results are obtained when demographic variables are included (column (b)).[25] This further suggests that the basic model fails to capture the effects of gender-specific confounding factors. Once again, the demographic variables appear to account for some of these but less adequately than the country

24. These are calculated as [exp (b_1LEAVE + b_2LEAVESQ)] − 1, for b_1 and b_2 the regression coefficients on the parental leave variables.

25. Female EP ratios are negatively related to birth and unemployment rates and positively correlated with employment shares in agriculture and services. Relative wages decline with birthrates and the share of agricultural employment but are unrelated to unemployment rates or the size of the service sector.

TABLE IV
DDD ESTIMATES OF THE EFFECTS OF PAID PARENTAL LEAVE USING QUADRATIC SPECIFICATIONS

Leave duration	Employment-to-population ratio (a)	(b)	(c)	(d)	Hourly wages (a)	(b)	(c)	(d)
Regression coefficients								
LEAVE	.0686	.1268	.2030	.2044	.0905	−.0922	−.1517	−.1350
	(.0806)	(.0710)	(.0593)	(.0867)	(.0672)	(.0611)	(.0532)	(.1100)
LEAVE SQUARED	−.0160	−.0647	−.2500	−.2521	−.1324	.2654	.2144	.1658
	(.0984)	(.0886)	(.0883)	(.1189)	(.0899)	(.0937)	(.1057)	(.1559)
p-value	.4616	.1071	.0034	.0628	.3281	.0096	.0167	.4704
Estimated differential versus no leave								
10 weeks	0.7%	1.2%	1.8%	1.8%	0.8%	−0.7%	−1.3%	−1.1%
20 weeks	1.3%	2.3%	3.1%	3.1%	1.3%	−0.8%	−2.2%	−2.0%
30 weeks	1.9%	3.3%	3.8%	3.9%	1.5%	−0.4%	−2.6%	−2.6%
40 weeks	2.5%	4.1%	4.2%	4.1%	1.5%	0.6%	−2.7%	−2.7%
Demographics	No	Yes	No	No	No	Yes	No	No
Time trends	No	No	Yes	Yes	No	No	Yes	Yes

See note to Table III. The dependent variables are differences in natural log of female and male labor market outcomes. LEAVE refers to the number of weeks of job-protected paid parental leave (irrespective of the wage replacement rate) divided by 100. The *p*-value refers to the null hypothesis that the coefficient on LEAVE and LEAVESQ are jointly equal to zero. All regressions include vectors of the country and year dummy variables. The lower panel of the table shows predicted differences in labor market outcomes, compared with no paid leave entitlement; these are calculated as $[\exp(b_1\text{LEAVE} + b_2\text{LEAVESQ})] - 1$, for b_1 and b_2 the regression coefficients on the parental leave variables. Specifications (a) through (c) are estimated using weighted least squares. OLS is used in specification (d), with Huber-White robust standard errors reported.

time trends. For this reason, the econometric models in the remainder of the paper include vectors of country-specific trends.[26] The results are insensitive to the method of accounting for heteroskedasticity, as evidenced by the virtually identical point estimates obtained in models (c) and (d). However, the Huber-

26. When demographic factors and country time trends are simultaneously controlled for, the results are close to those obtained with just time trends included (e.g., twenty weeks of paid leave increase predicted EP ratios by 3.8 percent and wages by 1.9 percent). The findings are also similar when time trends and unemployment rates (but no other demographic factors) are held constant. In this case, a twenty-week entitlement raises the expected EP ratio by 3.2 percent and lowers predicted hourly earnings by 2.1 percent. I also estimated models that included *lead* values of parental leave (at year $t + 1$), to provide a crude test of whether reverse causation is a problem. The leads of the leave variables did not jointly approach statistical significance (the *p*-value was .37 in the EP ratio equation and .85 for wages), and their inclusion had virtually no effect on the parameter estimates for contemporaneous leave in the wage models while modestly reducing the coefficients in the employment equations.

White standard errors are much larger than those obtained using the WLS procedure.

The estimates in Table IV suggest that parental leave mandates have large effects. For instance, rights to 40 weeks of job-protected paid leave are predicted to raise female EP ratios by 4.2 percent and lower hourly wages by 2.7 percent (in column (c)). As already mentioned, a portion of the positive relationship between leave durations and EP ratios may result because persons on parental leave are counted as "employed but absent from work," rather than "not employed." Some indication of the size of this effect at lengthier durations may be obtained by noting that 0.9 percent of the "employment" of 15–49 year old women in twelve EC countries was accounted for by maternity leave in 1983, as was 1.9 percent in 1992 [OECD 1995].[27] Since this rise in maternity leave usage occurred during a period when the average length of paid entitlements remained essentially constant, it may be reasonable to assume that around one percentage point of the increase in female EP ratios associated with extended paid entitlements is due to increased leave-taking, with two percentage points representing a likely upper bound.[28] Thus, this probably accounts for between one-quarter and one-half of the total estimated employment effect.

Table V provides information on the robustness of the results. Column (a) allows for lagged effects by including controls for job-protected leave (and its square) at year $t - 1$ as well as at time t. The table reports the sum of the coefficients, over the two years, along with the corresponding standard error; predicted effects in the bottom panel refer to changes in leave policies enacted at least one year previously. Column (b) controls for all types of paid leave, whether job-protected or not. The independent variable in column

27. The twelve countries include Belgium, Denmark, France, Germany, Greece, Ireland, Italy, Luxembourg, the Netherlands, Portugal, Spain, and the United Kingdom. An additional 0.6 percent of female employment in 1983 and 1.0 percent in 1992 was composed of work absences for personal or family reasons, a small portion of which may be due to parental leave. However, maternity leave will account for a smaller percentage of the employment of 15–64 year olds (the age group in this analysis) than of 15–49 year olds.

28. The population-weighted average duration of paid leave (full-pay weeks) declined from 34.1 (22.6) weeks in 1983 to 32.9 (22.0) weeks in 1992 for the eight countries (excluding Germany). There are three likely explanations for the increase in maternity leave over the period. First, some countries dramatically extended entitlements to unpaid leave, often with generous nonwork-related social insurance payments. Second, the persistence of high unemployment rates in many European nations may have made leave-taking relatively more attractive when compared with working. Third, increased labor force participation may have allowed women to meet the work history requirements needed to qualify for leave.

TABLE V
ADDITIONAL QUADRATIC SPECIFICATION ESTIMATES OF THE EFFECTS OF PAID PARENTAL LEAVE

Leave duration	Employment-to-population ratio (a)	(b)	(c)	(d)	Hourly wages (a)	(b)	(c)
Regression coefficients							
LEAVE	.2182	.0975	.3426	.2571	−.1694	−.2198	−.1944
	(.0667)	(.0984)	(.0692)	(.0664)	(.0609)	(.0913)	(.0604)
LEAVE SQUARED	−.2515	−.1322	−.6869	−.3662	.2428	.3110	.2676
	(.0965)	(.1337)	(.1333)	(.1314)	(.1192)	(.1550)	(.1225)
p-value	.0299	.6033	.0000	.0006	.0687	.0573	.0046
Estimated differential versus no leave							
10 weeks	1.9%	0.8%	2.8%	2.2%	−1.4%	−1.8%	−1.7%
20 weeks	3.4%	1.4%	4.2%	3.7%	−2.4%	−3.0%	−2.8%
30 weeks	4.4%	1.7%	4.2%	4.5%	−2.9%	−3.6%	−3.4%
40 weeks	4.8%	1.8%		4.5%	−2.8%	−3.4%	
Specification	Lagged leave	All paid leave	Full-pay weeks of leave	Balances sample	Lagged leave	All paid leave	Full-pay weeks of leave

See notes to Table III and IV. All regressions include country and year dummy variables, as well as country-specific time trends. Specification (a) controls for contemporaneous job-protected paid leave and parental leave lagged one year. The sum of the coefficients at t and $t-1$ are reported and the predicted effects refer to entitlements enacted at least one year previously. Specification (b) controls for all paid entitlements, whether or not job-protection is provided. In specification (c) the leave variable is full-pay weeks of leave, calculated as the weeks of job-protected paid leave multiplied by the estimated wage replacement rate. Specification (d) displays estimates for a sample restricted to observations where information on both employment-to-population ratios and hourly wages is available.

(c) is full-pay weeks of leave, defined as the average wage replacement rate multiplied by weeks of paid entitlement. In this case, the lower panel does not display an estimate for 40 weeks because fully paid entitlements of this duration are almost never observed in the sample. Finally, specification (d) displays results of an employment equation estimated for the eight countries (excluding Italy) for which wage data are available. These "balanced sample" estimates are directly comparable to those previously reported for wages.

Female EP ratios (wages) are positively (negatively) related to leave durations in all specifications. The estimated employment effects are larger for the "balanced sample" or when lags are included than the corresponding estimates in column (c) of Table IV, but the differences are modest, and the addition of lagged durations has virtually no effect on predicted wages. The coefficients on leave at $t - 1$ (not shown) suggest that changes in the entitlements have both an immediate impact and a considerably larger long-run effect, as anticipated if some laws are changed in the middle of the year or if adjustment to the new leave rights occurs over a period of time.[29]

Controlling for all types of paid leave, rather than just job-protected entitlements, results in smaller estimated employment effects and larger reductions in wages (see column (b)). For instance, 30 weeks of leave increases predicted female EP ratios by 1.7 percent and reduces expected wages by 3.6 percent, as compared with changes of 3.8 percent and −2.6 percent for rights to job-protected time off work. Countries instituting employment security provisions during the sample period initially moved from relatively short periods of nonprotected leave to similarly brief but protected entitlements. Therefore, the key issue determining whether the regressions should control for all leave or just job-protected absences is whether the addition of employment security provisions to existing leave rights (of short duration) has a substantial impact. As shown below, such changes are empirically important, particularly for EP ratios, implying that job-security provisions do need to be taken into account.

The wage replacement rate averages around 80 percent (85 percent for the first six months of leave and 72 percent thereafter), implying that a given duration of full-pay weeks would be expected to have roughly the same effect as a 25 percent longer entitlement to partially paid time off work. This is generally the case. For example, 30 full-pay weeks are predicted to increase female EP ratios by 4.2 percent and reduce wages by 3.4 percent (see column (c)), which is quite similar to the estimated effects previously obtained for 40 weeks of partially paid leave (4.2 percent and −2.7 percent).

B. "Step-Effects"

Continuous regressors may poorly capture the effects of the parental leave mandates, even when quadratic or higher order terms are included. For example, if women strongly wish to stay at home for a brief amount of time following childbirth, but with rapidly diminishing marginal utility of doing so thereafter, entitle-

29. The timing of the changes in parental leave policies is also frequently measured with an error of up to one year, which could induce the appearance of lagged effects and precludes a more sophisticated analysis of the dynamics of the adjustment process.

TABLE VI
ALTERNATIVE DDD SPECIFICATIONS EXAMINING THE EFFECTS OF JOB-PROTECTED PAID LEAVE

	Employment-to-population ratio			Hourly wages		
	(a)	(b)	(c)	(a)	(b)	(c)
Regression estimates						
ANYLEAVE	.0261	.0327	.0164	.0157	−.0109	.0078
	(.0137)	(.0084)	(.0104)	(.0119)	(.0070)	(.0090)
LEAVE	.0583		.2552	−.2507		−.2367
	(.0960)		(.0885)	(.0913)		(.0774)
LEAVESQ	−.0659		−.5634	.3647		.3259
	(.1302)		(.1541)	(.1541)		(.1395)
p-value	.0019	.0001	.0000-	.0189	.1254	.0088
Estimated differential versus no leave						
10 weeks	3.2%		3.7%	−0.6%		−1.3%
20 weeks	3.6%		4.6%	−2.0%		−2.6%
30 weeks	3.8%		4.3%	−2.6%		−3.3%
40 weeks	4.0%			−2.6%		

See the notes to Tables III through V. All specifications include country and year dummy variables and country-specific time trends. ANYLEAVE is a dummy variable taking the value one if the country has established an entitlement to job-protected paid leave and zero otherwise. LEAVE and LEAVESQ refer to weeks of paid job-protected leave (divided by 100) in column (a) and full-pay equivalent weeks of leave (weeks of job-protected paid leave entitlement multiplied by the average wage replacement rate) in specification (c). The *p*-value refers to the null hypothesis that the coefficients on ANYLEAVE, LEAVE, and LEAVESQ are jointly equal to zero.

ments to short job absences could have a substantial impact, whereas longer leave periods have little additional effect. The previously estimated models would then be misspecified since they restrict the consequences of rights to brief leaves to be small, relative to those of more extended durations. To permit this type of "step-effect," a dummy variable ANYLEAVE, which equals one if the country has enacted a leave mandate and zero otherwise, is included in Table VI. ANYLEAVE therefore indicates the impact of paid leave guarantees of arbitrarily short duration, with LEAVE and its square capturing the effects of extending an existing mandate. Job-protected paid leave is controlled for in the first two columns of the table; full-pay weeks are held constant in the third.

The coefficient on ANYLEAVE is of substantial size and has a *t*-statistic exceeding one in most specifications. The point estimates indicate that legislation requiring employers to offer minimal amounts of paid leave raise the relative employment of women by 1.7 to 2.6 percent and increase their hourly earnings by 0.8 to 1.6 percent (see specifications (a) and (c)). The predicted effect of further extending the leave period is summarized in the lower panel of the table.

These estimates confirm that paid leave is positively related to the EP ratios of women but suggest a more complicated story for wages. Whereas leave guarantees of substantial duration continue to be associated with sizable earning reductions, rights to brief periods away from the job now have less of an effect on expected wages. For instance, ten weeks of job-protected paid leave are predicted to reduce hourly earnings by 0.6 percent, as compared with a 1.3 percent reduction in the corresponding model that does not control for ANYLEAVE.

The results for all paid leave and full-pay weeks are also consistent. Since the average wage replacement rate is around 80 percent, we expect 10, 20, 30, and 40 weeks of job-protected paid leave to have approximately the same impact as 8, 16, 24, and 32 full-pay weeks. The estimated changes in EP ratios at these durations are 3.2, 3.6, 3.8, and 4.0 percent for (partially) paid leave versus 3.4, 4.3, 4.5, and 4.0 percent for full-pay weeks. The corresponding relative wage changes are −0.6, −2.0, −2.6, and −2.6 percent for job-protected entitlements, as compared with −0.9, −2.1, −3.0, and −3.4 percent for full-pay weeks.[30]

C. Women of Childbearing Age

I next compare the EP ratios of 25–34 year old women with those of same aged men or 45–54 year old females. (The lack of age-specific wage data precludes a similar analysis of earnings.) Since 25–34 year old females are in their prime childbearing years, they are expected to be strongly affected by parental leave mandates. Males of the same age constitute one possible comparison group; older women who have completed their fertility are

30. The robustness of the results was tested for by estimating specifications that included the ANYLEAVE dummy variable and a linear spline function. The results are similar to those reported. For example, when the break point is set at the median leave, conditional upon a positive entitlement (21 weeks for paid leave and 18 weeks for full-pay weeks), 10, 20, 30, and 40 weeks of job-protected paid leave raise predicted female EP ratios by 3.2, 3.6, 3.9, and 4.1 percent and decrease expected wages by 0.7, 2.2, 3.2, and 4.2 percent. Similarly, 10, 20, and 30 full-pay weeks raise employment by 3.0, 4.1, and 5.1 percent while cutting predicted wages by 0.9, 2.7, and 4.4 percent. Comparable results were obtained when setting the break points at 16 and 14 weeks, respectively (the twenty-fifth percentiles of nonzero entitlements).

TABLE VII
DDD ESTIMATES OF THE EFFECTS OF JOB-PROTECTED PAID PARENTAL LEAVE ON THE EP RATIOS OF WOMEN OF CHILDBEARING AGE

	Comparison group:			
	45–54 year old women		25–34 year old men	
	(a)	(b)	(a)	(b)
Regression estimates				
ANYLEAVE		.0465		.0285
		(.0259)		(.0238)
LEAVE	.0731	−.1559	.1291	−.0104
	(.1406)	(.1880)	(.1281)	(.1730)
LEAVESQ	.2313	.5076	.2366	.4052
	(.1748)	(.2309)	(.1594)	(.2124)
p-value	.0052	.0034	.0001	.0001
Estimated differential versus no leave				
10 weeks	1.0%	3.7%	1.5%	3.2%
20 weeks	2.4%	3.6%	3.6%	4.4%
30 weeks	4.4%	4.6%	6.2%	6.4%
40 weeks	6.8%	6.8%	9.4%	9.3%

See the notes to Tables III through VI. All specifications include country and year dummy variables and country-specific time trends. The dependent variable is the difference in the natural log of EP ratios between 25–34 year old women and either 45–54 year old women or 25–34 year old men. The sample contains 150 observations.

another.[31] Thus, the natural experiments in this section contrast changes in the percentage of the younger women employed to those of corresponding aged men or older females, as a function of variations in leave entitlements.

Table VII summarizes the econometric estimates. The dependent variable is the difference between (the log of) the EP ratios of 25–34 year old females and those of the comparison group. Time and country dummy variables are included, as are nation-specific time trends. Estimation is by weighted least squares. In addition to the quadratic in weeks of job-protected paid leave (divided by 100), specification (b) includes the ANYLEAVE dummy variable. Once again, the coefficient on ANYLEAVE is of substantial size

31. It is not obvious which comparison group is preferred. The use of younger men has the advantage of accounting for cohort differences that affect both sexes (such as the trend to extend education and delay entry into the labor force). However, changes in the employment of young females may have stronger effects on males of the same age, due to household labor supply decisions, than on older women.

and its inclusion improves the model fit, as measured by the adjusted R^2.

The table confirms that parental leave guarantees raise the employment of young women. The predicted changes in the EP ratios are larger than those for all females, as anticipated, since 25–34 year olds are in their prime childbearing years. For instance, entitlements to 40 weeks away from the job are predicted to increase the EP ratios of 25–34 year olds by 7 to 9 percent, compared with around 4 percent for all women.[32] However, these differences are less pronounced at short durations of leave.[33]

VII. Discussion

This analysis suggests that rights to short periods (three months) of paid parental leave increase the employment-to-population ratios of women by 3 to 4 percent while having little effect on wages. More extended entitlements (nine months) raise predicted female EP ratios by approximately 4 percent but decrease hourly earnings by around 3 percent. These effects are of similar magnitude to those obtained in some studies of other types of employment regulations and so need not be implausible.[34] Nevertheless, there are several reasons why they may overstate the true impact of parental leave guarantees. Most obviously, some countries may have provided additional rights to unpaid leave or implemented other "family-friendly" policies (such as subsidized child-care) at the same time they extended durations of paid time off work. In addition, if there are uncontrolled-for factors that simultaneously shift the female labor supply curve out and create political pressure to extend parental leave, lengthier leave would be correlated with higher EP ratios and reduced

32. The accounting bias resulting from counting persons on parental leave as "employed but absent from work" is likely to be larger for women of childbearing age than for all females. In 1983, 8.5 percent of "employed" women with children under five years of age, in twelve EC countries, were on maternity leave or were absent from work due to personal or family reasons; in 1992 the figure was 12.6 percent [OECD 1995]. Of course, many 25–34 year olds do not have children of this age.

33. Larger increases in employment are obtained in models that include the ANYLEAVE dummy variable and a linear spline. For instance, with a break point at 21 weeks, 30 weeks of parental leave are predicted to raise EP ratios of 25–34 year old women by 12 percent, compared with corresponding aged men, and by 14 percent compared with 45–54 year old females.

34. For example, Gruber [1994] estimates that mandated maternity benefits decrease the wages of some groups by up to 5 percent.

relative wages. Finally, it is possible that the entitlements encourage households to substitute female employment for male labor, violating the assumption that the legislated changes have no effect on the comparison group.

As mentioned, a positive relationship between leave durations and EP ratios may also occur because some individuals on parental leave are counted as "employed but absent from work" rather than "not employed." This cannot provide a full explanation for the findings, since a substantial rise in female EP ratios is predicted for rights to even short periods of paid leave, where the accounting bias should be relatively small. However, it may explain one-quarter to one-half of the increase in employment associated with longer entitlements.

Two primary factors are likely to account for the remaining 2 to 3 percent rise in women's employment. First, females who would not otherwise participate in the labor force may obtain jobs prior to childbirth in order to subsequently qualify for leave benefits. Some of the new entrants might also choose to remain employed after having children, but there is little evidence on the size of this effect, and it is not discussed further. Second, parental leave is likely to speed the return to work of new mothers.

The incentive for an individual to enter the labor force before having a child in order to qualify for parental leave is often strong. Persons working during the previous twelve months are generally eligible for leave benefits. Thus, a law establishing rights to three months of fully paid leave raises the effective wage for holding a job in the year prior to childbirth by 25 percent. Combined with substantial female labor force participation elasticities, this is likely to induce a substantial temporary increase in participation. Zabel [1996] estimates that the participation elasticities of women are between 0.5 and 1.0, implying that a 25 percent wage increase will raise female participation rates by 10 to 25 percent in the year before pregnancy.[35] With a baseline female EP ratio of 45 percent (the population-weighted sample average in these data), the enactment of three months of parental leave is therefore expected to induce a four to eleven percentage point increase in women's participation in the year before delivery which, if women average two children each, raises the average female EP ratio by between

35. Since the participation equation is nonlinear and the wage change is large, this should be viewed as a rough approximation.

0.4 and 1.0 percent.[36] Lengthier entitlements would presumably induce still larger entry into the labor force.

Previous research indicates that the availability of parental leave accelerates reentry into work. Klerman's [1993] analysis of the National Longitudinal Survey of Youth (NLSY) shows that the median time away from work for women receiving either paid or unpaid leave is around seven weeks. Conversely, the typical female who quits her job at or immediately before childbirth (presumably because she could not obtain sufficiently lengthy leave) does not return to work for more than one year. Waldfogel [1997] compares changes in work experience for NLSY women who do and do not return to the same employer after childbirth. The increase in experience, between the ages of 22 and 30, is 0.9 years greater for those who stay with the same employer than for those who do not (6.3 years versus 5.4 years).

Rönsen and Sundström [1996] study how parental leave affects the return to work in Norway and Sweden. Although both countries have relatively lengthy leave entitlements, those in Sweden are considerably more generous. It is presumably for this reason that Norwegian women are somewhat more likely to return to their jobs in the initial months following childbirth. However, Swedish reemployment rates ultimately considerably surpass those in Norway, with the result that young mothers in Sweden are almost twenty percentage points more likely to be employed three years after the first child is born.

These studies probably do not adequately control for all relevant sources of heterogeneity, so it remains uncertain to what extent leave mandates accelerate return to the labor force. Nevertheless, based on the available evidence, it may be reasonable to assume that entitlements of short-to-moderate length reduce the average time out of work by at least three months per child. If the typical woman has two children, this implies a six-month increase in job-holding which, with an average of around 23 years of lifetime employment, would raise the overall EP ratio by slightly over 2 percent. Longer leave periods may allow still more women to retain preexisting employment attachments but need not have a larger effect on female EP ratios, since

36. Female EP ratios average 45.3 percent, implying 1178 weeks of work between the ages of 15 and 64. Parental leave that induces a four percentage point increase in employment during the 52-week period preceding each of two births therefore raises the average EP ratio by 0.4 percent (4 percent $\times$ 104/1178). Similarly, a temporary eleven percentage point rise in employment elevates the EP ratio by 1.0 percent (11 percent $\times$ 104/1178).

they are likely to delay the return to work of individuals who would have otherwise done so fairly quickly.

Next consider wages. Entitlements to short periods of paid leave are estimated to have little impact on hourly earnings, whereas rights to lengthier time away from the job significantly reduce them. This seems reasonable. Brief periods of leave probably impose few costs on employers, particularly when the benefits are paid by the government. The entry of some new workers into the labor force will lower wages (as discussed below) but this may be more than offset by the increased experience associated with faster return to work following childbirth.

There are several reasons why rights to extended parental leave might substantially reduce wages. First, increases in labor supply in the period immediately prior to childbirth are likely to significantly lower female earnings. A small portion of this decline will result from the reductions in average experience levels.[37] It will mostly occur, however, because the outward shift of the labor supply curve is combined with inelastic demand for labor. Hamermesh [1993] estimates that the long-run own-price elasticity of demand for homogeneous labor is around -0.3, implying that a 1 percent rise in labor supply will reduce wages by more than 3 percent. The decrease in women's relative wages will be less than this, to the extent that male and female labor is easily substitutable, but Hamermesh suggests that such substitutability is limited.

Second, extended work absences may impose substantial nonwage costs on firms. As the leave entitlement lengthens, it is likely to become much harder to schedule replacement workers, particularly in countries placing restrictions on the duration of employment on temporary fixed-term contracts. The costs may be sizable even when workers do not use all of the allowed leave, since employers face considerable uncertainty regarding the timing and ultimate likelihood of the individual's return to the job.

Third, extremely lengthy leave guarantees introduce the possibility that women having multiple children in a short period of time may be away from their jobs for several years consecutively, or with just brief spells of intervening employment, causing substantial depreciation of human capital. This becomes even

37. If the leave entitlements induce a 1 percent rise in the female EP ratios and inexperienced workers earn 20 percent less than the average women, the compositional change will decrease female wages by less than 0.2 percent.

more likely when paid leave is supplemented by rights to unpaid but job-protected time off work.

To summarize, this study indicates that parental leave guarantees raise the employment of women but, at longer durations, may be paid for through the receipt of lower relative wages. Several mechanisms have been identified as possible sources of the relatively large estimated effects. Nevertheless, beyond the caveats already mentioned, these findings should be viewed as tentative for a variety of reasons. The sample sizes are quite small, resulting in imprecise estimates in some models. The data on leave are incomplete—a more comprehensive investigation would explicitly consider eligibility and take-up rates. And the information on wages is often restricted to manufacturing workers. Ideally, this analysis would be supplemented by research using microdata from several countries. Finally, other benefits or costs may also be associated with the mandates. Most significantly, it is often argued that parental leave improves the health and well-being of children. This represents an important area for future research.

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