PARENTAL LEAVE AND CHILD HEALTH

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Working Paper **6554**

NBER WORKING PAPER SERIES

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Working Paper 6554
http://www.nber.org/papers/w6554

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
May 1998

I thank Robert Clark, Janet Currie, Jane Waldfogel, and workshop participants at North Carolina State University, the National Bureau of Economic Research, University of New Hampshire, and the American Economic Association Meetings for helpful comments on earlier versions of this paper. This work has been supported by a grant from the National Institute on Aging to the National Bureau of Economic Research. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

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NBER Working Paper No. 6554
May 1998
JEL Nos. I12, I18, J38

ABSTRACT

This study investigates whether rights to paid parental leave improve pediatric health, as measured by birth weights and infant or child mortality. Aggregate data are used for nine European countries over the 1969 through 1994 period. Year and country fixed-effects are held constant and most specifications include additional covariates or control for country-specific time trends. Much of the analysis incorporates a "natural experiment" comparing changes in pediatric outcomes to those of senior citizens, whose health is not expected to be affected by parental leave. More generous leave rights are found to reduce deaths of infants and young children. The magnitudes of the estimated effects are substantial, especially for those outcomes where a causal effect of parental leave is most plausible. In particular, there is a much stronger negative relationship between leave durations and post-neonatal mortality or fatalities between the first and fifth birthday than for perinatal mortality, neonatal deaths, or the incidence of low birth weight. The evidence further suggests that parental leave may be a cost-effective method of bettering child health.

Christopher J. Ruhm
Department of Economics
University of North Carolina Greensboro
P.O. Box 26165
Greensboro, NC 27402-6165
and NBER
c_ruhm@uncg.edu

Parental Leave and Child Health

Over 100 countries, including virtually all industrialized nations, have enacted some form of parental leave policies (Kamerman, 1991). Most assure women the right to at least two or three months of paid leave during the period surrounding childbirth. Proponents believe these entitlements improve the health of children and the position of women in the workplace. Opponents counter that the mandates reduce economic efficiency, by restricting voluntary exchange between employers and employees, and may have particularly adverse effects on the labor market opportunities of females. These disagreements persist, in part, because the results of requiring employers to provide parental leave are poorly understood.

A small but rapidly growing literature has examined the effects of these policies on labor market outcomes.[1] By contrast, to my knowledge, only two studies provide any information on the relationship between parental leave and health. First, using data for 17 OECD countries, Winegarden and Bracy (1995) find that an extra week of paid maternity leave correlates with a two to three percent reduction infant mortality rates. The accuracy of these results is questionable, however, because the estimated effects are implausibly large and are sensitive to the treatment of wage replacement rates during the job absence – for example, short or medium durations of leave at high replacement rates are projected to increase infant deaths in some specifications. The lack of robustness may be due to small sample sizes or limitations in the

[1] Analysis of the U.S. for the period before the enactment of federal legislation generally finds that time off work is associated with increases in women's earnings and employment (e.g. Dalto, 1989; Spalter-Roth and Hartmann, 1990; Waldfogel, 1997). However, this may result from nonrandom selection into jobs providing the benefit, rather than the leave itself. Recent studies attempt to overcome the selection problem by focusing on state regulations (Kallman, 1996; Klerman and Leibowitz, 1997), federal legislation (Waldfogel, 1996), or leave mandates in Europe (Ruhm and Teague, 1997; Ruhm, 1998). Results of this research are mixed. The preponderance of evidence suggests that leave rights increase female employment but possibly

methodological approach and imply that the findings should be interpreted cautiously.[2] Second, McGovern, et al. (1997) indicate that time off work has nonlinear effects on the postpartum health of mothers, as measured by mental health, vitality, and role function. Specifically, short-to-moderate periods away from the job (up to 12 to 20 weeks) are associated with worse health, whereas the reverse is true for longer absences. This pattern is difficult to explain using any plausible health production function and probably does not show a causal effect. Instead, it is likely that the quadratic specification used is overly restrictive, that a nonrandom sample of women take time off work after birth, or both.

This study provides the most detailed investigation to date of the relationship between parental leave entitlements and pediatric health. Aggregate data are used for nine European countries over the 1969 through 1994 period.[3] The outcomes examined are the incidence of low birth weight and several types of infant and child mortality. Time and country effects are controlled for and additional covariates or country-specific time trends are often included to capture the effects of confounding factors that vary over time within countries.

Even with the aforementioned controls, it is possible that omitted determinants of health (such as lifestyles or medical technologies) are correlated with parental leave. One method of accounting for these effects is identify a comparison group whose health is affected by the same unobservables as that of young children, but for whom parental leave is anticipated to have little

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with an accompanying decline in relative wages for lengthy entitlements.

[2] The estimating equation has fewer than 70 observations and 50 degrees of freedom. In addition: the fixed-effect models employed are unlikely to adequately account for time-varying confounding factors; the definition of paid leave probably includes payments that are independent of previous employment histories; and the equations do not allow for nonlinear effects of leave durations or replacement rates.

[3] A distinction is sometimes made between "maternity leave", granted to mothers for a limited period around the time of childbirth, and "parental leave" which permits additional time off to

impact. Senior citizens are chosen for this purpose and many of the results are for a "natural experiment" examining how changes in leave rights affect the difference between mortality rates of the young and old.

Understanding the effects of parental leave is important for both Europe and the United States. Europe has been struggling with the question of whether social protections inhibit economic flexibility and employment growth (Blank, 1994; Siebert, 1997; Nickell, 1997). All Western European countries currently offer at least three months of paid maternity benefits but many of the policies have been instituted or significantly revised during the period analyzed, and some nations have recently shortened the leave period or reduced the payments provided during it (Organization for Economic Cooperation and Development, 1995). By contrast, the United States did not require employers to provide parental leave until the 1993 enactment of the Family and Medical Leave Act (FMLA), and advocates (e.g. the Carnegie Task Force on Meeting the Needs of Young Children, 1994) have argued for broadening the law to cover small establishments and provide payment during the work absence.[4]

To preview the results, rights to parental leave are associated with substantial decreases in pediatric mortality, especially for those outcomes where a causal effect is most plausible. In particular, there is a much stronger negative relationship between leave durations and either post-neonatal mortality (deaths between 28 days and one year of age) or child fatalities (deaths

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care for infants or young children. Both are included in the definition of parental leave used here.

[4] The FMLA requires employers with more than 50 workers in a 75-mile area to allow 12 weeks of unpaid leave to persons with qualifying employment histories following the birth of a child or for a variety of health problems. There are exemptions for small firms and certain highly paid workers. A number of states enacted limited rights to leave prior to the FMLA and many workers could also take time off work under the provisions of the Pregnancy Discrimination Act of 1978 or by using vacation or sick leave. See Ruhm (1997a) for further discussion of the provisions and effects of the FMLA.

between the first and fifth birthday) than for perinatal mortality (fetal deaths and deaths in the first week), neonatal mortality (deaths in the first 27 days), or the incidence of low birth weight. The evidence further suggests that parental leave may be a cost-effective method of bettering child health.

1. Parental Leave and the Health of Children

The health of young children depends on many factors including: the "stock" of health capital, the level of medical technology, the price of and access to health care, household income, and the time investments of parents. As discussed below, parental leave is most likely to improve pediatric health through the last of these mechanisms.[5]

The stock of health capital is partially stochastic but also depends on previous investments and lifestyle choices (Grossman, 1972).[6] However, most of these investments occur early in pregnancy and so are unlikely to be substantially enhanced by European leave policies, which generally provide time off work for only a short period immediately prior to birth (usually six weeks).[7] There could even be negative effects. Specifically, the availability of paid leave may induce some women to work early in their pregnancies, in order to meet the employment requirements to qualify for it. This reduces the time available for health investments (such as

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[5] These reduced-form relationships can be derived from the maximizing the utility function $U(H,X)$, subject to the budget constraint $Y=P_mM+P_xX=wR+sL+N$, the time constraint $T=R+L+V$, and the health production function $H(B,M,L+V,\varepsilon)$. H, X, M, and Y are health of the child, other consumption, medical care, and total income. P_m and P_x are relative prices; T, R, L, and V indicate total time, time at work, time on leave, and nonmarket time. B is baseline health, ε a stochastic shock, w the wage rate, s the payment during parental leave, and N is nonearned income. Time away from work (L+V) is assumed to be positively related to children's health.

[6] For example, smoking or drinking by pregnant women may impair fetal development and result in high rates of low weight births, perinatal deaths, and neonatal mortality (Chomitz, Cheung, and Lieberman, 1995; Frisbie, Forbes, and Pullman, 1996).

[7] Modest benefits are possible. For instance, parental leave may facilitate bed-rest late in pregnancy, where indicated to reduce the probability of premature birth, and some countries

early prenatal care) and could lead to higher rates of still births and mortality during the first months of life.[8]

Medical care can raise the stock of health capital. Intensive interventions (e.g. neonatal intensive care) are particularly crucial for remedying deficits during the early days of life and are associated with substantial reductions in neonatal mortality (Corman and Grossman, 1985; Currie and Gruber, 1997). The medical infrastructure and most lifestyle choices are unlikely to be affected by parental leave entitlements but may be correlated with them, and so need to be controlled for in the analysis.

Higher incomes may improve health by raising access to medical care, particularly when a substantial portion of the expenditures are paid out-of-pocket, and by increasing the purchase of other health-improving goods and services (e.g. diet, sanitation, safety).[9] Rights to parental leave are likely to modestly elevate the percentage of women employed and, unless fully offset by reductions in wages or spousal labor supply, raise household incomes. However, the increase is probably quite small and so the effect on pediatric health is likely to be minimal.[10]

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require employers to permit lengthier absences before birth if there is a medical reason to do so.

[8] The induced employment may be substantial. Ruhm (1998) estimates that a law establishing three months of fully paid leave will increase female labor supply by 10 to 25 percent in the year before pregnancy. Women in industrialized countries almost always obtain prenatal care prior to childbirth; however, many do not receive it sufficiently early in their pregnancy. Studies examining the determinants of birth weights or fetal and neonatal mortality therefore typically focus on whether care is provided in the first trimester, or on the number of months from the beginning of pregnancy until prenatal care is first received (e.g. Rosenzweig and Schultz, 1983; Grossman and Joyce, 1990; Frank, et. al, 1992; or Warner, 1995).

[9] However, the relationship between income and health is ambiguous for industrialized countries. Some studies uncover a positive association (e.g. Ettner, 1996) while others find no effect (e.g. Duleep, 1995). Ruhm (1997b) shows that many types of health are adversely affected by short-lasting improvements in economic conditions, with less negative or more beneficial effects for sustained economic growth. There is stronger evidence that incomes and health are positively related in developing countries (e.g. see Pritchett and Summers, 1996).

[10] Ruhm (1998) estimates that rights to substantial leave induce a 3 to 4 percent increase in

Parental leave is likely to primarily affect child health by making more time available to parents. As recognized by Becker (1981, Chapter 5), raising children is an extremely time-intensive activity. The commitments begin before birth (e.g. the need for greater sleep and for adequate prenatal care) but are likely to be particularly large during the first months of life. Moreover, some important time investments present special logistical challenges for employed persons, and so may be facilitated by rights to leave.

Breast-feeding is an example of one such activity. The consumption of human milk by infants is linked to better health through decreased incidence or severity of many diseases (e.g. diarrhea, lower respiratory infection, lymphoma, otitis media, and chronic digestive diseases), reductions in infant mortality from a variety of causes (including sudden infant death syndrome), and possibly enhanced cognitive development.[11] However, it is often more difficult for working women to breast-feed and employment reduces both its frequency and duration (Ryan and Martinez, 1989; Gielen et al., 1991; Lindberg, 1996; Blau, et al., 1996; Roe, et al., 1997).

Many health ailments afflicting the very young are transitory and have little impact on long-term development. From a policy perspective, the greatest concern is for problems that have lasting effects and, in the extreme, result in death.[12] For this reason, mortality rates are the primary proxy for health in the analysis below. One way to conceptualize the relationship between mortality and health is to define a minimum threshold level of health capital, H_{min},

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female employment. This probably represents an upper bound on the rise in household income because many new mothers have working spouses or receive income from social insurance or other sources. Kallman (1996) and Ruhm also provide evidence of partially offsetting reductions in female wages.

[11] See Cunningham, Jellifee, and Jelliffee (1991) or the American Academy of Pediatrics (AAP) Work Group on Breast-feeding (1997) for reviews of the benefits of breast-feeding. The AAP has recently recommended that infants be fed human milk for the first twelve months of life.

[12] Of course, even relatively minor illnesses can escalate into fatal health problems.

below which death occurs. The expected level of health H^* is a function of the various inputs into the health production function and realized health is defined by $H = H^* + \varepsilon$, where ε is a stochastic shock. The probability of death is:

$$\text{Pr (Mortality)} = \Pr(\varepsilon \leq H_{min} - H^*) = \Phi(H_{min} - H^*), \tag{1}$$

where $\Phi(.)$ is the c.d.f. of the error term. Mortality and health are therefore inversely related and are affected by (many of) the same determinants.

2. Data

The analysis uses annual aggregate data covering the years 1969 through 1994 for nine nations: Denmark, Finland, France, (the Federal Republic of) Germany, Greece, Ireland, Italy, Norway, and Sweden. The countries selected are all Western European nations with substantial changes in leave policies during the sample period.[13] In order to distinguish between job-related leave and social insurance payments that are independent of previous work histories, paid leave is defined to include rights to job absences where the size of the income support depends on prior employment. Most of the investigation focuses on job-protected leave, where dismissal is prohibited during pregnancy and job-reinstatement is guaranteed at the end of the leave.[14] A

[13] Gaps and noncomparabilities in the data become much more severe prior to 1969 and parental leave policies changed little during the early and middle 1960s. Other European nations (Austria, Belgium, the Netherlands, Portugal, Spain, Switzerland, and the United Kingdom) were excluded because their leave policies did not change significantly during the time span of analysis.

[14] Until recently, women were generally *prohibited* from working during specified periods surrounding childbirth and frequently received neither income support nor guarantees of job-reinstatement. Starting in the late 1960s, maternity leave began to evolve to emphasize paid and job-protected time off work, with father's increasingly gaining rights to leave. However, vestiges of protective legislation persist, with postnatal leaves remaining compulsory in many nations and prenatal leave continuing to be required in some. See Organization of Economic Cooperation and Development (1995), Ruhm and Teague (1997), or Ruhm (1998) for additional discussion of the history of European leave policies.

measure of "full-pay" weeks is also calculated by multiplying the duration of the leave by the average wage replacement rate received during it.

The leave entitlements 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 have not changed or have loosened over time in most countries, however, 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 and the secular increase in parental leave entitlements is probably understated.[15]

Unpaid leave has not been incorporated into this analysis for two reasons. First, many employers may be willing to provide time off work without pay, even in the absence of a mandate, making it difficult to distinguish between the effects of job absences voluntarily granted by companies 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. Also, no attempt is made to distinguish leave available only to the mother from that which can be taken by either parent, or to model differences in "take-up" rates. These restrictions should be kept in mind when interpreting the results. If within-country growth in paid entitlements is positively correlated with changes in the proportion of persons with qualifying work histories or rights to unpaid leave, the econometric estimates will combine these factors and may overstate the impact of an increase in paid leave that occurs in isolation.

[15] This discussion focuses on women because they take the vast majority (usually far above 95 percent) of total weeks of parental leave, even when the rights extend to fathers.

In 1986, Germany simultaneously lengthened the duration of job-protected leave and extended to nonworkers the income support previously restricted to persons meeting qualifying employment conditions (Ondrich, et al. 1996). Using the above criteria, this would be defined as a reduction in paid leave (since payments were no longer tied to prior employment). However, such a classification seems problematic, since the duration of job-protected time off work was substantially increased in 1986 and again in later years. For this reason, data for Germany are included only through 1985.[16]

Information on parental leave is 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*, published biennially by the United States Social Security Administration.[17] The wage replacement rates used to calculate full-pay weeks of leave are approximations because they do not account for minimum or maximum payments and because some nations provide a "flat rate" amount or a fixed payment plus a percentage of earnings.[18]

Table 1 summarizes parental leave provisions in effect during the last year of the data (1994 except for Germany). At that time, the 9 countries offered a minimum of 14 weeks of paid leave and 6 nations provided rights to more than 6 months off work. Full-pay weeks ranged from 9 weeks in Greece to 58 weeks in Sweden, with a positive correlation between replacement rates

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[16] As an alternative, models were estimated with German leave entitlements either assumed to remain constant (at 32 weeks) after 1985, or increasing according to the extensions granted in subsequent years. In the first case, the estimated parental leave effects are virtually identical to those detailed below. In the second, the predicted decreases in mortality tended to be somewhat larger (smaller) in models that include (exclude) controls for fertility and EP ratios.
[17] This is an updated version of the parental leave data in Ruhm (1998). A subset of the information was also used by Teague (1993) and Ruhm and Teague (1997). Jackqueline Teague played a primary role in the initial data collection effort.
[18] In these cases, the replacement rate is estimated as a function of average female wages, using data from various issues of the International Labor Office's *Yearbook of Labour Statistics*. See

and leave durations. Income support 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 2 displays leave durations and estimated wage replacement rates for each country at five year intervals. The number of nations providing job-protected leave rose from 4 in 1969 to 6 in 1979, with all 9 doing so after 1983. Countries supplying parental benefits in 1969 extended them subsequently, with the result that the dispersion of leave entitlements tended to increase over time. There were 30 observed changes in durations over the sample period and 5 additional cases where nations modified replacement rates without altering the length of leave.

Pediatric health is proxied in the analysis by the incidence of low birth weight and several mortality rates. Information on birth weight and perinatal deaths is obtained from the OECD *Health Data 96* (Organization for Economic Cooperation and Development, 1996). Data on neonatal, post-neonatal, infant, child, and senior citizen mortality are from the WHO *Health for All Data Base: European Region* (World Health Organization, 1997).[19] Table 3 provides definitions and descriptive statistics for all variables used below.

Data limitations restrict the set of regressors included in the econometric models. The characteristics sometimes controlled for include: real per capita GDP (GDP), health care expenditures as a percent of GDP (SPENDING), the share of the population with health insurance coverage (COVERAGE), the number of kidney dialysis patients per 100,000

Ruhm (1998) for details.

[19] In the WHO data, child mortality refers to deaths before age 5. This was converted into deaths between the first and fifth birthday by subtracting infant mortality rates.

population (DIALYSIS), the fertility rate of 15-44 year old women (FERTILITY), and the female employment-to-population ratio (EP RATIO).[20]

GDP, SPENDING, COVERAGE, and DIALYSIS are expected to be positively related to child health. Higher incomes allow greater investments in health capital and medical care. Holding income constant, health is likely to improve when a greater proportion of spending is for medical care and when health insurance is common. The use of kidney dialysis is *not* anticipated to be causally related to pediatric outcomes. Rather, it proxies sophisticated medical technologies (e.g. neonatal intensive care) for which data are not available.

Female employment could affect pediatric health by changing income and nonmarket time. For instance, working women may have less time to invest in infants, leading to worse health. Similarly, several studies (e.g. Rosenzweig and Wolpin, 1988; Frank, et al., 1992) suggest that fertility rates and infant deaths are positively correlated. However, these variables may be endogenous, since parental leave is often provided or extended with the goal of increasing birthrates or improving the labor market opportunities of women.[21] There may also be reverse causation. Higher infant mortality rates imply, *ceteris paribus*, that more births are needed to achieve a target family size and that there are fewer young mothers, who have

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[20] Data are from Organization for Economic Cooperation and Development (1996). Several procedures were used to fill in missing values for some variables. In particular: 1969 values for DIALYSIS were extrapolated assuming a constant growth rate between 1969 and 1971; FERTILITY for France and Denmark in 1969 was assumed to be the same as in 1970; and French fertility in 1971-1974 was interpolated using a linear trend between 1970 and 1975. EP RATIOS are from Ruhm (1998), with values in the early years for Greece and Norway set equal to those in the first period for which data were available (1972 and 1977 respectively).
[21] Averett and Whittington's (1997) analysis of U.S. data indicates women working for employers providing maternity leave have modestly higher fertility rates than those who do not.

relatively low rates of employment.[22] Reflecting these concerns, results will be presented for models both with and without these "supplemental" regressors.

The econometric techniques are designed to account for omitted factors and cross-country differences in the definition or measurement of the dependent variables.[23] In addition to the limited set of covariates, the models control for general time effects, country fixed-effects, and (frequently) country-specific time trends. Also, senior citizens are often included as a comparison group. The implicit assumption is that the death rate of persons 65 and over (DEATH 65) is affected by the same unobserved factors (e.g. changes in lifestyle or medical technologies) as pediatric outcomes, but are not causally affected by parental leave.

3. Time Trends

Parental leave entitlements rose sharply between 1969 and 1994. Weighting by the number of births in each cell, the mean duration of job-protected paid leave grew from 11 to 32 weeks and full-pay weeks from 9 to 25 weeks (see Figure 1a).[24] The growth was most dramatic prior to 1980, with a particularly large jump occurring at the end of the 1970s when 6 countries (Finland, France, Germany, Italy, Norway, and Sweden) almost simultaneously extended entitlements. Since 1980, there has been little change in average 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 weeks were provided at relatively low replacement rates.

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[22] See Browning (1992) for discussion of the relationship between children and female labor supply.

[23] Liu, et al. (1992) document significant cross-national differences in the measurement of infant mortality.

[24] These calculations assume that German parental leave entitlements remained constant at 32 weeks after 1985.

Figure 1b documents trends in the child health outcomes. Observations are displayed as percentages of 1969 values (1970 for child mortality) and are weighted by the number of births. There is no evidence that the incidence of low birth weight has fallen over time. The instability observed early in the period occurs because data are missing for several countries during some of these years.[25] Nevertheless, even after the middle 1980s, when the information becomes more complete, there is no indication of a downward trend.[26] This is not surprising. Birth weight results from a complex interaction of factors. For instance, improvements in prenatal care probably raise birth weights but this may be offset by new medical technologies that increase the survival of low-weight fetuses. Thus, birth weight provides an ambiguous measure of pediatric health and strong associations between it and parental leave are unlikely.

By contrast, pediatric mortality has fallen dramatically since the late 1960s. Infant fatalities decreased 75 percent between 1969 and 1994 (from 23.1 to 5.7 per thousand live births), perinatal deaths by 76 percent (from 27.0 to 6.6 per thousand live and still births), and child mortality by 68 percent between 1970 and 1994 (from 3.5 to 1.2 per thousand live births).[27] Obviously, most of these reductions are unrelated to parental leave, highlighting the importance of controlling for potential sources of spurious correlation.

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[25] For instance, there is a spike in 1973 because this is the only year prior to 1979 that data are available for Italy (which has relatively high rates of low-weight births). Birth weight information is also missing in several years for France, Greece and Ireland.

[26] This mirrors the experience of the United States, where the incidence of low weight births *rose* modestly between the middle 1980s and early 1990s.

[27] Neonatal deaths fell 79 percent (from 16.1 to 3.4 per 1000 live births) and post-neonatal mortality by 64 percent (from 7.0 to 2.5 per 1000 thousand live births) between 1969 and 1994. By comparison, the standardized death rate of senior citizens fell 37 percent (from 68.4 to 43.0 per 1000 population) over the period.

4. Estimation Strategy

Health outcome H for group i in country j at year t is assumed to be determined by:

$$H_{ijt} = a_1G_i + a_2C_j + a_3T_t + b_1(G_i \times C_j) + b_2(G_i \times T_t) + b_3(C_j \times T_t) + d_iL_{jt} + e_{ijt}. \quad (2)$$

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

A model that controls for time and country fixed-effects therefore estimates:

$$H_{ijt} = \alpha + \beta_1C_j + \beta_2T_t + \delta L_{jt} + \varepsilon_{ijt}, \quad (3)$$

where $\alpha = a_1G_i$, $\beta_1 = a_2 + b_1G_i$, $\beta_2 = a_3 + b_2G_i$, $\delta = d_i$ and $\varepsilon_{ijt} = b_3(C_j \times T_t) + e_{ijt}$. This is a "difference-in-difference" (DD) model, since it examines within-country growth in the dependent variable as a function of modifications in leave durations. However, equation (3) does not contrast the changes to those of a comparison group expected to be unaffected by the leave entitlements (the third difference in the DDD model below). 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 increase entitlements at a time when health is improving for other reasons (e.g. due to new medical technologies). Notice, however, that a direct indication of this bias can be obtained from the estimated value $\hat{\delta}$ for an outcome (such as mortality of the elderly) not expected to be affected by parental leave.

Letting the subscripts c and o indicate children and older persons, the difference in health outcomes between these two groups can be expressed as:

$$H_{cjt} - H_{ojt} = a_1(G_c - G_o) + b_1(G_c - G_o)C_j + b_2(G_c - G_o)T_t + (d_c - d_o)L_{jt} + (e_{cjt} - e_{ojt}) \quad (4)$$

or equivalently

$$\Delta H_{jt} = \alpha + \beta_1 C_j + \beta_2 T_t + \delta L_{jt} + \varepsilon_{jt}. \tag{5}$$

Equation (5) is a "difference-in-difference-in-difference" (DDD) model. β_1 and β_2 indicate group-specific country and time differences, δ shows the age-difference in the impact of parental leave, and, importantly, differencing has eliminated the time-specific country effect (C_j x T_t) from the regression error term. Thus, these estimates measure how growth in the age disparity in health outcomes varies as a function of within-country changes in leave entitlements.

If parental leave has no effect on the health of senior citizens, d_o=0 and the DDD estimate $\hat{\delta}$ supplies an unbiased estimate of d_c. By contrast, if $d_o \neq 0$, and has the same sign as d_c, the regression coefficient will be biased towards zero. This might occur if parental leave improves the health of older persons by increasing the time available for their adult children to assist them. However, the resulting bias is likely to be small, since leave rights are restricted to the period surrounding childbirth, whereas the elderly will frequently need help at other times.

Equation (5) accounts for time-varying factors that affect the two groups equally. However, the estimates will still be inconsistent if within-country changes in parental leave are correlated with unobservables (e.g. the availability of high-quality day care) that have different effects on the health of the young and old. This can be seen by adding a third level interaction $c_1(G_i$ x C_j x $T_t)$ to equation (2). The error term in (5) becomes $\varepsilon_{ijt} = c_1(G_c - G_o)(C_j \text{ x } T_t) + (e_{cjt} - e_{ojt})$, which may be correlated with L_{jt}. Omitted explanatory variables represent a potentially important source of group-specific time-varying factors. A vector of country-specific linear time trends will therefore frequently be included in the models, since many of the unobserved factors (e.g. the level of medical technology) are likely to exhibit a monotonic trend.[28]

5. Results

The econometric results are summarized in this section. The relationship between leave entitlements and death rates of senior citizens is examined first. As discussed, parental leave is not expected to have much impact on health of the elderly, so any observed association probably results from a failure to control for omitted factors. Thus, these regressions guide the choice of subsequent specifications and indicate if there is likely to be a gain from including a comparison group in the examination of pediatric health. Next, a detailed investigation is provided of the determinants of infant mortality, a common measure of child health. This is followed by consideration of the other outcomes – low birth weight and perinatal, neonatal, post-neonatal or child mortality. Finally, the estimating equations are modified to allow nonlinearities.

Vectors of country and time dummy variables are included in all models, additional covariates or country-specific time trends are frequently controlled for, and the dependent variables are measured in natural logs. The outcomes in the DDD equations are differences between the (natural logs of the) relevant child health measure and the death rate of senior citizens. Weighted least squares (WLS) is used to adjust for heteroskedasticity resulting from differences in cell sizes. The weights are determined using a three-step procedure. First, the equations are estimated by OLS. Next, the squared residuals from these models are regressed against a constant and the reciprocal of the number of births in the cell. The square root of the

[28] The omitted variable problem is even more severe in the DD models, since these do not include a comparison group. For example, bias will be introduced by almost any changes in medical technology that are correlated with parental leave entitlements, whether or not they disproportionately benefit the young, making it even more crucial to include country-specific

inverse predicted values from the second-stage regression are then used as weights in the final estimates. Blackburn [1997] shows that this procedure is more efficient than weighting by (the square root of) cell sizes or using OLS with robust standard errors if there is a common group effect or group-time interaction across individuals in a country.[29]

5.1 Mortality Rates of Senior Citizens

Table 4 summarizes econometric estimates of equations where the dependent variable is the death rate of persons aged 65 and over. Parental leave is measured as a quadratic in weeks of job-protected paid entitlement divided by 100 and the p-value refers to the null hypothesis of no relationship between leave rights and mortality. The leave coefficients are jointly significant (at the .06 level or higher) in all specifications that exclude country-specific time trends (see columns a through c), indicating the need to be concerned about confounding variables in the examination below. However, when country time-trends are included there is no longer a statistically significant relationship (column d), suggesting that these do a better job of controlling for omitted factors than the available national characteristics.

The other regressors generally have the expected signs. Mortality rates are negatively related to per capita GDP, health care spending as a percent of GDP, and the level of medical technology (as proxied by the prevalence of dialysis). The positive effect of insurance coverage is surprising but does not show up in the children's health equations discussed below. The

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trends in these specifications.

[29] A large and significant constant term is obtained in virtually all of the second-stage regressions, confirming the importance of group effects and justifying the use of this procedure. Indeed, the coefficient on the reciprocal of the number of births is usually small and *insignificant*, suggesting that unweighted estimates would be preferable to using the standard method of weighting by the square root of cell sizes.

coefficient on fertility is small and insignificant but a strong positive effect is observed for female EP ratios, possibly because employed women have less time to assist unhealthy parents.

5.2 Infant Mortality

Table 5 displays the results of eight specifications examining the determinants of infant mortality. The parental leave regressor is weeks of job-protected paid leave divided by 100. The dependent variable in the DDD models is the difference in the (natural log of) mortality rates of infants and senior citizens. Additional covariates and country-specific time trends are included in some specifications. As expected, higher income, greater health spending, broader insurance coverage, and increased medical technology (indicated by the frequency of dialysis) reduce predicted infant mortality rates, usually by statistically significant amounts.[30]

Parental leave is estimated to have a substantial negative effect on infant mortality. For instance, a 10 week extension in leave is predicted to decrease infant deaths by 1.7 percent in column (a) and 1.6 percent in model (e). As discussed, country-specific time trends probably provide a preferable method of accounting for omitted factors. Thus, it is informative that the parental leave coefficient rises (in absolute value) when these are controlled for – rights to 10 extra weeks reduce predicted mortality by 2.5 percent in specification (c) and 2.1 percent in column (g).[31] The estimated impact becomes greater still when fertility rates and female EP

[30] Specifications equivalent to columns (b) and (f), but with country time trends included, were also estimated. GDP, SPENDING, COVERAGE, and DIALYSIS were neither individually nor jointly significant in these models and so the results are not displayed.

[31] I tested the sensitivity of the results to the choice of weighting procedures by reestimating columns (c) and (g) using OLS, but with Huber-White robust standard errors, and alternatively using the square root of the number of births as cell weights. The parameter estimates were virtually identical to those reported in the first case (-.2531 and -.2153), although the standard errors were larger (.2210 and .2081). In the second case, the coefficients were only 50 to 60 percent as large, although still marginally significant. However, as mentioned, the first-stage regressions strongly reject the appropriateness of this weighting procedure.

ratios, hereafter referred to as "supplemental regressors", are held constant – the decrease in expected deaths is 2.6 percent in column (d) and 2.5 percent in model (h). In large part, this occurs because female employment is strongly positively related to both infant mortality rates and parental leave durations.

Table 6 tests the robustness of the findings to a variety of alternative specifications. This table and the remainder of the analysis focuses on equations that include country-specific time trends. The first row restates the results obtained in columns (c), (d), (g), and (h) of Table 5. The second panel refers to models that include lagged, as well as current, leave entitlements. The third controls for all paid leave, whether or not job-protection is provided. The fourth holds constant full-pay weeks of leave.

The negative relationship between parental leave and infant mortality persists across specifications. There appears to be a stronger long-run than short-run effect, as evidenced by the negative coefficient on lagged leave rights. This is logical since the regulations sometime change in the middle of a calendar year, the leave period may span years, and the adjustment to any policy change may be gradual.[32] The estimated impact of all paid leave (with or without job-protection) or full-pay weeks is marginally larger than that of job-protected leave. The magnitudes also tend to be somewhat greater in the DD than DDD specifications and when the supplemental regressors are controlled for, but the differences are not large or statistically significant. More impressive is the overall consistency of results – a 10 week increase in leave entitlements reduces predicted infant mortality rates by between 2 and 3 percent in all models.

[32] I also estimated models with *lead* values of leave included, as a crude check of reverse causation. The lead coefficient was always small and never approached statistical significance.

5.3 Other Health Outcomes

The results for other pediatric outcomes, summarized in Table 7, are entirely consistent with those expected if parental leave has a *causal* effect on children's health. There is no evidence that leave improves fetal development, as measured by birth weight or perinatal mortality. This is anticipated since the time off work generally occurs late in the pregnancy and employment may be induced during its early stages. Similarly, the small effect on neonatal mortality is logical given that deaths in the first month of life are primarily determined by health at birth and medical interventions during the period surrounding it.[33]

Conversely, leave entitlements substantially reduce predicted mortality during the post-neonatal period and in early childhood – a 10 week extension decreases expected post-neonatal deaths by 4.5 to 6.6 percent and child fatalities by 2.6 to 3.1 percent. Benefits during the post-neonatal period make sense, since this is when parental leave is most likely to result in additional time at home with children. There may also be longer-lasting gains since leaves sometimes extend beyond one year and investments made during the first 12 months may yield future health benefits.

Fertility rates (female EP ratios) are negatively (strongly positively) correlated with post-neonatal fatalities. The DDD coefficients (not displayed) imply that an increase in the fertility rate from 1.8 to 2.0 children reduces predicted post-neonatal mortality by 6.2 percent, while a 10 percentage point decrease in the percentage of women employed does so by 14.9 percent. Both effects are highly significant. The fertility result may reflect economies of scale in raising

[33] Also, mothers can often take time off work during the first month after birth, even without leave rights. For example, 73 percent of "employed" women with one-month-old infants were on leave (41 percent on paid leave) rather than working in the U.S. during the 1986-1988 period, prior to the passage of federal parental leave legislation (Klerman and Leibowitz, 1994).

children. The employment finding is consistent with the possibility that working mothers have less time to invest in their offspring.[34]

5.4 Nonlinearities

There are several reasons why the relationship between parental leave and the pediatric health may be nonlinear. First, the proportion of the entitlement actually used may vary with its length. For example, some persons may not be able to afford extended leaves with partial wage replacement. Second, the marginal benefit of time investments in infants may decline with their age. Either factor will induce diminishing returns. Conversely, workers may be able to leave their jobs for short but not long periods, in the absence of a formal mandate, implying that legislation providing brief leaves will have no effect on health, whereas benefits will be obtained from lengthier durations.

The form of the nonlinearity may also vary across outcomes. For instance, neonatal mortality is unlikely to be reduced by extensions of postnatal leaves beyond one month. Since maternity leave rights typically begin (in Europe) around six weeks before birth, this implies that there should be little marginal benefit to leave durations exceeding 10 weeks. Conversely, short

[34] The employment effect is *not* fully explained by reverse causation, whereby higher infant death rates reduce the number of new mothers, who work relatively infrequently. In this case there would also be a strong positive association between EP ratios and neonatal mortality rates, which we do not observe – a ten percentage point reduction in female employment is predicted to raise neonatal mortality by a statistically insignificant 2.8 percent in the DDD model. There could be a spurious correlation between female employment and fatality rates. To examine this possibility, I reestimated the models with both female and *male* EP ratios controlled for. The results were mixed. The positive effect of female employment on post-neonatal mortality became even stronger and the male coefficient was insignificantly negative. However, female (male) EP ratios were significantly negatively (positively) related to perinatal and neonatal mortality. These ambiguous findings suggest that more analysis is needed to determine how parental employment affects child health.

entitlements could speed the return to work and so raise post-neonatal and possibly child mortality, whereas lengthier leave periods are expected to reduce these sources of death.[35]

Nonlinearities are modeled by linear spline specifications with knots at 13, 26, and 39 weeks of leave. Table 8 and Figures 2a and 2b display DDD estimates of changes in predicted mortality at various leave durations, compared to the case of no entitlement.[36] The first p-value on the table refers to the null hypothesis of no parental leave effect. The second tests whether the inclusion of the splines significantly improves model fit. The supplemental regressors are included as additional covariates in column (b) of Table 8 and in Figure 2b.

There is strong evidence of nonlinearities, as indicated by the joint significance of the splines, and the results are once again consistent with those expected if parental leave has a causal effect on health. In particular, reductions in predicted perinatal and neonatal mortality are modest (although statistically significant) and concentrated on rights to time off work in the period immediately surrounding birth. By contrast, extended entitlements sharply decrease post-neonatal and child mortality rates, whereas rights to brief leave either have no effect or slightly increase them – 20 weeks of parental leave increase expected deaths during these periods by 1 and 3 percent respectively, whereas a 50 week entitlement is predicted to reduce post-neonatal fatalities by at least 20 percent and child mortality by 11 to 12 percent.[37]

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[35] The reason that rights to short leave may hasten the return to work is that individuals must choose between a brief absence, but with the right to continue in the original position, and a longer time off but with eventual reemployment in a new (and probably less desirable) job. In some cases, it will be worthwhile to accept a shorter leave period so as to avoid the change of employers. See Klerman and Leibowitz (1997) for further discussion.

[36] The DD results are similar but typically indicate slightly larger benefits of leave. There is never a significant effect on low birth weight, so these findings are not displayed.

[37] Similar results were obtained when nonlinearities were modeled by including polynomials in leave. For example, with a cubic specification and no supplemental regressors, 10, 20, 30, 40, and 50 weeks of leave reduce predicted post-neonatal mortality by -3.5, -1.2, 4.8, 12.2, and 19.3

The reduction in post-neonatal mortality associated with lengthy leave is larger when the supplemental regressors are controlled for than when they are not – rights to 50 weeks of leave decrease the expected death rate by 29 percent in the former specification versus 20 percent in the latter. Much of the disparity occurs because female employment is positively correlated with both post-neonatal fatalities and leave entitlements. To the extent that the increases in job-holding are exogenous, the estimates obtained in the absence of controls for them therefore represent a lower-bound on the parental leave effect. Conversely, if extensions in leave cause female employment to increase, the results obtained with the EP ratio held constant will overstate the true impact. A reasonable guess is that the actual effect of parental leave is approximately half-way between the two estimates.[38]

6. Plausibility and Cost-Effectiveness

The econometric estimates suggest that parental leave entitlements substantially reduce mortality during early childhood – rights to a year of job-protected paid leave are associated with roughly a 25 percent decline in post-neonatal deaths and an 11 percent decrease in fatalities occurring between the first and fifth birthdays. Effects of these magnitudes are large but not unreasonable. Post-neonatal mortality and child mortality fell more than 60 percent during the

percent. This compares to -2.3, -1.3, 5.0, 15.6, and 20.1 percent in the model with splines.

[38] Ruhm (1998) estimates that a year of parental leave rights raises female employment by around 4 percent. This corresponds to a 1.8 percentage point rise, at the sample average EP ratio of 45.1 percent. The regression coefficient on the female EP ratio is 1.628 (with a t-statistic of 3.7) in the post-neonatal mortality equation, implying that this increase in employment will cause post-neonatal deaths to rise by 3 percent ($e^{(.018 \times 1.628)} - 1 = .0297$). Subtracting this from the upper-bound estimate, implies that 50 weeks of parental leave reduces post-neonatal fatalities by around 26 percent, after accounting for the additional employment it induces.

sample period, implying that the decreases predicted to result from extensions in leave rights are small compared to those that actually transpired.[39]

Moreover, there are a variety of mechanisms through which parental leave might yield substantial health benefits. As mentioned, time off work may increase breast-feeding. Roe, et al. (1997) estimate that an extra week of postpartum job absence raises the duration of breast-feeding by 3 to 4 days, with an accompanying growth in frequency for those who do so. Although it is difficult to determine the extent to which this might reduce infant deaths, the available evidence suggests the effect could be substantial. For example, a 30 percentage point increase in the fraction of women intending to breast-feed was estimated to decrease post-perinatal death rates by more than 9 percent, after controlling for a other risk factors, in Carpenter, et al.'s (1983) analysis of a prevention program in Sheffield England. Similarly, Cunningham, et al. (1991) find that breast-feeding is associated with a 3.7 per 1000 fall in post-perinatal mortality, although some of this may be due to omitted factors. Based on these results, a reasonable guess is that a substantial parental leave entitlement might increase breast-feeding sufficiently to prevent 0.5 to 1.0 post-neonatal deaths per 1000 live births. This represents a 6 to 13 percent reduction in this source of mortality, compared to the 1969 sample average.

Parental leave may also reduce a number of specific risks of death during early childhood. To illustrate, Table 9 summarizes the leading causes of infant and young child mortality in the United States.[40] Not surprisingly, neonatal fatalities are dominated by health problems

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[39] Average leave durations grew from 11 weeks in 1969 to 32 weeks in 1994. A rise of this size is predicted to reduce post-neonatal mortality by 8 percent and child fatalities by 1 to 4 percent.
[40] Data for neonatal and post-neonatal mortality are for 1992; those for child mortality are for 1995. More recent information (which is not yet available for neonatal or post-neonatal deaths) is provided in the last case because it breaks out fatalities resulting from HIV, whereas the 1992 data do not. However, the rankings and percentages of child deaths from different causes are

originating during the prenatal or birth periods. By contrast, parental inputs are likely to play a major role in deterring many subsequent deaths. For example, four of the five leading causes of post-neonatal mortality (Sudden Infant Death syndrome, accidents, pneumonia/influenza, and homicide), accounting for 48 percent of fatalities, are almost certainly substantially influenced by the activities of parents.[41] Similarly, accidents and homicides account for 43 percent of child fatalities and several other leading causes (e.g. heart disease, HIV, pneumonia/influenza) may be sensitive to parental involvement.

An obvious policy question is whether the health benefits of parental leave are worth the costs. Towards this end, the appendix summarizes estimates of the government expenditure on parental leave payments required to save one child's life.[42] The key assumptions are that: 1) one week of parental leave entitlement causes a .0000295 reduction in the probability of death; 2) each week of leave rights translates into between .18 and .34 weeks of actual time away from work; 3) annual earnings during the leave period average $22,000.

Using these assumptions, the cost per life saved is between $2.6 and $4.9 million (in $1997). These amounts are within the general range of estimates typically obtained from value-

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virtually identical in 1992 and 1995.

[41] Closer parental involvement is likely to prevent some accidental deaths and may indirectly reduce other sources of fatalities. For example, Sudden Infant Death Syndrome is more than twice as common among infants who sleep prone as for those who do not (Hunt, 1996; Taylor et al., 1996; Øyen et al., 1997). Parental leave could increase the frequency of non-prone sleeping if parents have more energy to monitor sleeping position or are more able to directly observe it. Time off work might also decrease homicides by reducing stress levels in families with young children. Finally, parental leave might lessen the need for child care, which is associated with increased risk of many infectious illnesses (e.g. see Redmond and Pichichero, 1984; Thacker et al., 1992; or Hardy and Fowler, 1993). Parental inputs may even influence mortality due congenital anomalies to the extent they determine whether the child receives timely medical treatment and other health-preserving investments.

[42] Government expenditures are used as the measure of cost, since leave payments are received from the government in all sample nations.

of-life calculations, suggesting that the provision of parental leave may be a cost-effective method of improving health. For example, Viscusi (1992, p. 73) states that most "reasonable estimates of the value of life are clustered in the \$3 to \$7 million range", and Manning et al., (1989) use a figure of \$1.66 million. Adjusting for inflation (using the all-items CPI), these are equivalent to \$3.5 to \$8 million and \$2.15 million, respectively, in 1997 dollars,

Moreover, several factors will lead these this analysis to understate the net benefits of parental leave. First, the measured health improvements are limited to reductions in mortality, whereas many gains may take the form of better health for living children. Second, there may be advantages for children and families other than health (e.g. improved cognition or reductions in household stress). Third, previous research suggests that leave rights may improve the labor market status of women. Fourth, the leave payments may partially offset other types of government spending (e.g. by reducing the utilization of subsidized child care or decreasing public spending on medical services), lowering the true cost of providing it.

Of course, there are many uncertainties associated with the calculations and the analysis that underlies them, some of which could lead to overly favorable assessments. For example, the range of value of life estimates is quite large, with lower valuations sometimes placed on children than adults (since human capital investments have not yet been made) and on individuals in low-income households.[43] The sample sizes are also quite small, resulting in imprecise estimates in some specifications, and neither eligibility for nor take-up of parental leave has been explicitly modeled. Most importantly, some sources of confounding may not have been fully accounted for. In particular, some countries may have provided unpaid leave or

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[43] Currie and Gruber (1996) provide a careful discussion of these issues, in the context of reductions in infant mortality resulting from Medicaid expansions in the United States.

implemented other "family-friendly" policies (such as subsidies for high-quality child-care) at the same time they extended durations of paid time off work, so that the measured effects of the latter actually capture the combined impact of a variety of factors.

Although concerns about spurious correlation are legitimate, it is unlikely that they are a source of significant bias. As discussed, the econometric estimates rely on within-country variations, after controlling for linear time trends and factors that have similar effects on the health of the young and old. Substantial bias would therefore be caused by discrete changes in policies or environmental factors that are focused on young children and occur at virtually the same time as modifications in parental leave legislation. The most likely candidates are child and health care policies. Compared to the United States, European countries often do have more systematic and generous protections in both these areas. However, there is considerable heterogeneity across nations, including substantial cross-sectional variation among countries with relatively parental leave entitlements, and little reason to believe that modifications in these policies are closely correlated with changes in parental leave.[44] Moreover, these sources of confounding would probably lead to a correlation between leave rights and perinatal or neonatal

[44] For example, during the late 1980s, 44 percent of Danish children aged 3 and below received child care outside the home, compared to just 2 percent in Germany and 5 percent in Italy, even though all three countries offered more than 6 months of paid leave (Kamerman, 1991). By contrast, France offered limited paid leave (16 weeks) but used day care relatively extensively (24 percent of 0-3 year olds) and, as documented by Bergmann (1996), supplied pregnant women and young children with comprehensive and targeted medical services. Similarly, Kamerman and Kahn (1993) provide evidence that the diverse use of home health visiting (HHV) services in European countries is relatively unrelated to the provision of paid leave. For instance, they claim that Great Britain and Denmark have the most comprehensive HHV programs, whereas other nations (e.g. Finland, Germany, Italy, Norway, and Sweden) offer more extensive leave entitlements. Further evidence of the heterogeneity of policies aimed at young children is provided by Kamerman and Kahn (1991), Gustafson and Stafford (1994), or Organization for Economic Cooperation and Development (1995).

mortality. Instead, the association is restricted to the post-neonatal and young child periods, where parental leave is most likely to have a causal effect.[45]

7. Conclusion

This analysis lends considerable credence to the view that parental leave has favorable and possibly cost-effective impacts on pediatric health. The most likely reason is that the work absences provide parents with additional time to invest in their young children. This may be increasingly important since female labor force participation rates have risen substantially over time in most countries and given the evidence that women's employment is positively correlated with infant mortality. More generally, the results suggest that time is an important but poorly understood input into the production of health. Ideally, future research will verify the results of this study using microdata and making a special effort made to identify the mechanisms through which leave entitlements improve health.

[45] For instance, Finland provides extensive health benefits through special maternal and child health clinics. To receive maternity benefits and parental allowances, pregnant women and their babies are required to undergo regular prenatal and postnatal care – women average 18 prenatal visits per pregnancy and infants visit the clinics 13 times during the first year of life (Mikkola, 1991). Since Finland also provides extended parental leave, the health and leave benefits could be correlated. However, if omitted controls for health care were the cause of the "parental leave" effects, we would expect to see an impact on fetal development or health outcomes during early infancy, which is largely absent from the data.

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Appendix: Estimates of the Cost-Per-Live Saved From Parental Leave

Entitlements to 1 week of (additional) paid leave are assumed to have no effect on neonatal mortality but to decrease post-neonatal and child mortality by .56% and .28% respectively.[46] At the sample averages of 4.204 and 2.124 deaths per 1000 live births, this implies a total reduction of .0295 fatalities per 1000 births, or a .0000295 decrease in the probability of death. Two estimates are provided of amount of leave actually taken by mothers. The first "high" estimate assumes that 42.5 percent of pregnant women work throughout their pregnancies and that these individuals use 80 percent of available leave benefits, implying .34 weeks of actual work absence (.8 x .425 = .34) per week of entitlement.[47] These probabilities are based on Swedish data and almost certainly overstate the use of leave because Sweden has relatively high female labor force participation rates prior to birth and "take-up" rates subsequent to it. A second, and probably preferable, "low" estimate assumes that 30 percent of pregnant women are employed throughout pregnancy and that this group uses 60 percent of available leave benefits.[48] This implies that a week of leave rights results in .18 weeks of actual use (.6 x .3 = .18).

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[46] These are mid-points of the DDD estimates, with and without supplemental regressors, detailed in Table 7.

[47] Sundstöm and Stafford (1992) estimate that 48 (8) percent of women worked full-time (part-time) immediately before giving birth to their first child and 20 (26) percent prior to later children during the 1988-90 period. Assuming that part-time workers were employed half as many hours as those employed full-time and that half of infants were first births (the total fertility rate was 1.96 in 1988), "full-time equivalent" employment was 42.5% (0.5(.48 + .08/2) + 0.5(.20 + .26/2)). Rönsen and Sundstöm (1996) estimate that 81 percent of married mothers in Sweden, whose first birth occurred after age 18, used parental leave.

[48] Klerman (1993) estimates that around 70 percent of new mothers in the United States, which also has relatively high female participation rates, either do not work at all during pregnancy or quit their jobs prior to giving birth.

The "high" estimate implies that 11,525 weeks (.34/.0000295) or 221.6 years of parental leave are required to save one life. Using the "low" estimate, 6,102 weeks (.18/.0000295) or 117.3 years are needed. If mothers employed full-time are paid $11 per hour for 2000 hours per year, their annual earnings are $22,000.[49] This suggests a cost-per-life saved of 4.9 million dollars (221.6 x $22,000 = $4,875,200) using the "high" estimate and 2.6 million dollars (117.3 x $22,000 = $2,580,600) by the "low" estimate.

[49] Waldfogel (1997) indicates that the average hourly earnings of 30 year old women in the U.S. were $9.40 in 1991. Adjusting for price changes, using the all-items CPI, this was equivalent to $11.08 per hour in 1997. These estimates could be adjusted upwards or downwards in a variety of ways. For instance, some fringe benefit costs could be added. On the other hand, wage replacement rates average around 80 percent during the leave period. These effects are likely to roughly cancel each other out.

Table 1: Job-Protected Paid Parental Leave in 1994

Country	Leave Entitlement	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.

Note: Information for Germany refers to 1985.

Table 2: Job-Protected Paid Parental Leave and Wage Replacement Rates in Selected Years

Country	1969	1974	1979	1984	1989	1994
Denmark	0	0	0	18 [.90]	28 [.90]	28 [.90]
Finland	0	29 [.55]	35 [.55]	43 [.80]	44 [.80]	44 [.80]
France	0	0	16 [.90]	16 [.90]	16 [.90]	16 [.84]
Germany	14 [1.00]	14 [1.00]	32 [1.00]	32 [1.00]		
Greece	0	0	0	12 [.60]	12 [.60]	15 [.60]
Ireland	0	0	0	14 [.70]	14 [.70]	14 [.70]
Italy	21 [.80]	31 [.80]	57 [.57]	48 [.53]	48 [.53]	48 [.53]
Norway	12 [.13]	12 [.32]	18 [1.00]	18 [1.00]	24 [1.00]	42 [1.00]
Sweden	16 [.55]	26 [.90]	39 [.90]	52 [.71]	52 [.71]	64 [.90]

Note: The wage replacement rates, which are shown in brackets, are sometimes subject to minimum or maximum amounts. Also, they are sometimes estimated to account for differing replacement rates during early and later portions of the leave or flat rate payments.

Table 3: Summary Information on Variables Used in Analysis

Variable	Definition and Descriptive Statistics
	Outcome Variables (per 1,000 live births unless noted)
INFANT	Infant Mortality: infant deaths under 1 year ($n=232$, $\mu=13.1$, $\sigma=6.6$)
LOW WEIGHT	Low Birth Weight: new-borns weighing less than 2,500 grams as % of live births and still births over 1,000 grams ($n=172$, $\mu=5.5$, $\sigma=.63$)
PERINATAL	Perinatal Mortality: stillbirths (≥ 28 weeks gestation) and deaths within 1 week of birth per 1,000 live & still births ($n=228$, $\mu=15.5$, $\sigma=7.3$)
NEONATAL	Neonatal Mortality: infant deaths under 28 days ($n=220$, $\mu=8.9$, $\sigma=5.3$)
POSTNEO	Post-neonatal Mortality: deaths between 28 days and 1 year ($n=220$, $\mu=4.2$, $\sigma=1.7$)
CHILD	Child Mortality: deaths between 1 and 5 years of age ($n=223$, $\mu=2.1$, $\sigma=.75$)
DEATH65	Standardized Death Rate of Persons ≥ 65 years old per 1000 population ($n=233$, $\mu=55.9$, $\sigma=8.7$)
	Other Variables
LEAVE	Weeks of Job-Protected Paid Parental Leave ($n=225$, $\mu=23.5$, $\sigma=17.2$)
PAID	Weeks of Paid Parental Leave with or without job-protection ($n=225$, $\mu=25.6$, $\sigma=14.8$)
RATE	Average wage replacement rate (in %) during Parental Leave ($n=225$, $\mu=79.6$, $\sigma=17.5$)
GDP	Real GDP per capita in thousands of 1994 U.S. dollars, adjusted using PPP and the all-items CPI ($n=234$, $\mu=16.0$, $\sigma=3.0$)
SPENDING	Expenditures on Health Care as Percent of GDP ($n=234$, $\mu=7.3$, $\sigma=1.5$)
COVERAGE	Share of population with Health Insurance coverage ($n=234$, $\mu=.954$, $\sigma=.057$)
DIALYSIS	Number of Dialysis patients per 100,000 population ($n=234$, $\mu=15.8$, $\sigma=11.2$)
FERTILITY	Fertility Rate of 15-44 year old women ($n=233$, $\mu=1.81$, $\sigma=.42$)
EP RATIO	Female Employment-to-Population Ratio: civilian employment divided by the 15 to 64 year old population, using standardized OECD definitions ($n=230$, $\mu=.451$, $\sigma=.107$)
BIRTHS	Number of Births in thousands ($n=234$, $\mu=614$, $\sigma=266$)

Note: Observations are weighted by the number of births in each cell.

Table 4: Econometric Estimates of the Mortality Rates of Persons Aged 65 and Older

Regressor	(a)	(b)	(c)	(d)
LEAVE	.2406 (.0955)	.1668 (.0898)	.1212 (.0850)	-.1189 (.0716)
LEAVE SQUARED	-.4361 (.1222)	-.2748 (.1146)	-.3178 (.1097)	.1466 (.1106)
GDP		-.0053 (.0041)	-.0134 (.0043)	
SPENDING		-.0205 (.0053)	-.0157 (.0051)	
COVERAGE		.0021 (.0007)	.0026 (.0007)	
DIALYSIS		-.0022 (.0008)	-.0017 (.0008)	
FERTILITY			.0064 (.0121)	
EP RATIO			.5415 (.1085)	
p-value	.0016	.0588	.0069	.2514
Time Trends	No	No	No	Yes

Note: The dependent variable is the natural log of the standardized death rate of persons aged 65 and over. Data are for nine European countries over the 1969-1994 period (n=225). Standard errors are shown in parentheses. All models include country and year dummy variables. The estimates in column (d) also include country-specific (linear) time trends. Observations are weighted to correct for heteroscedasticity using the procedure discussed in the text. LEAVE refers to weeks of job-protected parental leave divided by 100. The p-value refers to the null hypothesis that the coefficients on LEAVE and LEAVE SQUARED are jointly equal to zero.

Table 5: Econometric Estimates of the Effects of Paid Parental Leave On Infant Mortality Using Linear Specifications

Regressor	DD Estimates				DDD Estimates			
	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)
LEAVE	-.1703	-.4013	-.2538	-.2623	-.1633	-.3275	-.2160	-.2517
	(.1169)	(.1053)	(.0854)	(.0864)	(.1108)	(.1048)	(.0851)	(.0847)
GDP	-.0223	-.0284			-.0172	-.0156		
	(.0090)	(.0086)			(.0085)	(.0086)		
SPENDING	-.0233	-.0140			-.0027	.0020		
	(.0116)	(.0102)			(.0110)	(.0102)		
COVERAGE	-.0095	-.0056			-.0121	-.0085		
	(.0014)	(.0014)			(.0014)	(.0014)		
DIALYSIS	-.0089	-.0073			-.0065	-.0055		
	(.0018)	(.0015)			(.0017)	(.0015)		
FERTILITY		.1328		.0445		.1349		.0165
		(.0243)		(.0356)		(.0243)		(.0349)
EP RATIO		1.044		.3460		.5081		.6599
		(.2163)		(.2206)		(.2154)		(.2161)
Time Trends	No	No	Yes	Yes	No	No	Yes	Yes

Note: See note on Table 4. All equations include vectors of year and country dummy variables (n=223). The dependent variable in the DD estimates is the natural log of the infant mortality rate. In the DDD estimates, the dependent variable is difference between (the natural logs of) the infant mortality rate and the standardized death rate of persons aged 65 and over.

Table 6: Alternative Linear Specifications Examining the Effects of Paid Leave on Infant Mortality

Parental Leave Regressor	DD Estimates (a)	DD Estimates (b)	DDD Estimates (c)	DDD Estimates (d)
Job-Protected Paid Leave	-.2538 (.0854)	-.2623 (.0864)	-.2160 (.0851)	-.2517 (.0847)
Job-Protected Paid Leave in Current Year	-.0895 (.1292)	-.0943 (.1293)	-.0541 (.1290)	-.0814 (.1269)
Job-Protected Paid Leave in Previous Year (t-1)	-.2220 (.1317)	-.2280 (.1312)	-.2188 (.1315)	-.2311 (.1288)
All Paid Leave	-.2642 (.0958)	-.2975 (.0967)	-.2272 (.0954)	-.2841 (.0950)
Full-Pay Weeks of Leave	-.2797 (.1098)	-.2842 (.1103)	-.2775 (.1087)	-.3081 (.1076)
Supplemental Regressors	No	Yes	No	Yes

Note: See notes on tables 4 and 5. Each panel refers to a separate series of regressions. All equations include year and country dummy variables, as well as country-specific time trends. "All Paid Leave" refers to paid entitlements, whether or not job-protection is provided. "Full-Pay Weeks of Leave" is calculated as the number of weeks of job-protected leave multiplied by the estimated wage replacement rate. Weeks of parental leave are divided by 100 throughout the table. "Supplemental Regressors" include the fertility rate and female employment-to-population ratio.

Table 7: Econometric Estimates of the Effects of Job-Protected Paid Parental Leave On Various Pediatric Outcomes Using Linear Specifications

Health Outcome	DD Estimates		DDD Estimates	
	(a)	(b)	(c)	(d)
Low Birth Weight	-.0777 (.0937)	-.1232 (.0965)	-.0075 (.0974)	-.0464 (.1015)
Perinatal Mortality	-.0365 (.0838)	.0548 (.0773)	.0075 (.0765)	.0682 (.0733)
Neonatal Mortality	-.0601 (.1163)	-.0033 (.1143)	-.0132 (.1090)	-.0141 (.1081)
Post-Neonatal Mortality	-.5044 (.1617)	-.6881 (.1477)	-.4553 (.1714)	-.6699 (.1538)
Child Mortality	-.3181 (.1375)	-.2679 (.1372)	-.2989 (.1334)	-.2732 (.1338)
Supplemental Regressors	No	Yes	No	Yes

Note: See notes on tables 4 through 6. Each panel refers to a separate series of regressions. All equations include year and country dummy variables, as well as country-specific time trends. The coefficients displayed are for weeks of job-protected paid leave divided by 100. Low Birth Weight refers to new-borns weighing less than 2,500 grams, perinatal mortality to stillbirths and deaths within the first week of life, neonatal mortality to deaths in the first 27 days, post-neonatal mortality to those occurring between days 28 and 365, and child mortality indicates fatalities between the first and fifth birthday. Sample sizes are 163, 219, 214, 211, and 211 for the five outcomes.

Table 8: Linear Spline Estimates of the Percentage Reductions in Mortality Due to Job-Protected Paid Parental Leave

Weeks of Leave	Infant Mortality		Perinatal Mortality		Neonatal Mortality		Post-Neonatal Mortality		Child Mortality	
	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)	(a)	(b)
10	2.3	1.5	3.6	1.0	4.0	1.0	-2.3	3.7	-1.0	-3.8
20	2.5	3.0	4.2	1.9	4.4	2.8	-1.3	5.3	-3.1	-5.0
30	5.4	7.6	4.0	2.3	4.8	4.8	5.0	12.0	0.2	0.0
40	12.5	15.2	4.1	1.7	6.5	6.2	15.6	25.2	10.7	10.6
50	10.5	12.1	0.4	-2.8	1.7	0.0	20.1	29.1	12.4	10.9
p-value: leave	.0055	.0001	.0307	.0459	.1123	.0621	.0046	.0000	.0054	.0021
p-value: splines	.0391	.0025	.0140	.0318	.0588	.0304	.0435	.0448	.0203	.0053

Note: The table displays the predicted percentage reduction in mortality associated with the specified weeks of job-protected paid leave, compared to no leave mandate. The estimates are obtained from DDD models that include controls for country and year effects, as well as country-specific time trends. Specification (b) includes the fertility rate and female employment-to-population ratio as supplemental regressors, whereas column (a) does not. The linear splines are estimated with knots at 13, 26, and 39 weeks. The first p-value refers to the null hypothesis that parental leave has no effect on the outcome; the second refers to the null hypothesis that parental leave is linearly related to the dependent variable (i.e. no splines are needed).

Table 9: Leading Causes of Neonatal, Post-neonatal, and Child Deaths

Cause of Death	# of Deaths	% of Deaths
Neonatal Mortality		
1. Congenital anomalies (740-759)	5375	24.6
2. Disorders relating to short gestation/unspecified low birthweight (765)	3971	18.2
3. Respiratory distress syndrome (769)	1927	8.8
4. Maternal complications of pregnancy (761)	1456	6.7
5. Complications of placenta, cord, membranes (762)	975	4.5
6. Infections specific to perinatal period (771)	837	3.8
7. Intrauterine hypoxia and birth asphxia (778)	543	2.5
8. Neonatal hemorrhage (772)	307	1.4
9. Sudden infant death syndrome (798.0)	298	1.4
10. Birth trauma (767)	199	0.9
Post-Neonatal Mortality		
1. Sudden infant death syndrome (798.0)	4593	35.9
2. Congenital anomalies (740-759)	2074	16.2
3. Accidents (E800-E949)	736	5.8
4. Pneumonia and influenza (480-487)	478	3.7
5. Homicide (E960-E969)	278	2.2
6. Septicemia (038)	225	1.8
7. Respiratory distress syndrome (769)	136	1.1
8. Bronchitis and bronchiolitis (466,490-491)	121	.9
9. Malignant neoplasms (140-208)	98	.8
10. Meningitis (320-322)	91	.7
Child Mortality		
1. Accidents (E800-E949)	2280	35.7
2. Congenital anomalies (740-759)	695	10.9
3. Malignant neoplasms (140-208)	488	7.6
4. Homicide (E960-E969)	452	7.0
5. Heart disease (390-398,402,404-429)	251	3.9
6. Human immunodeficiency virus (042-044)	210	3.3
7. Pneumonia and influenza (480-487)	156	2.3
8. Certain conditions originating in the perinatal period (760-779)	87	1.4
9. Septicemia (038)	80	1.3
10. Cerebrovascular diseases (430-438)	57	0.9

Note: Data for neonatal and post-neonatal deaths are for 1992 (source: National Center for Health Statistics, 1996). Information on child mortality is for 1995 (source: Anderson, Kochanek, and Murphy, 1997). ICD-9 codes are in parentheses.

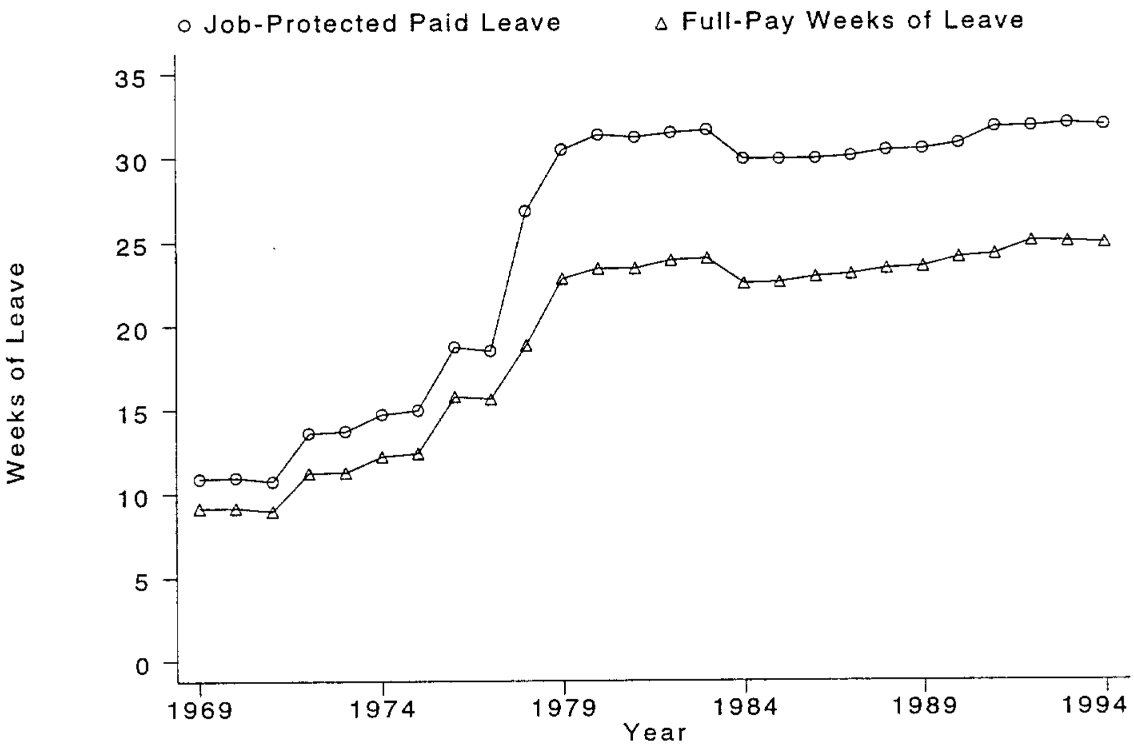

Figure 1a: Average Weeks of Paid Parental Leave

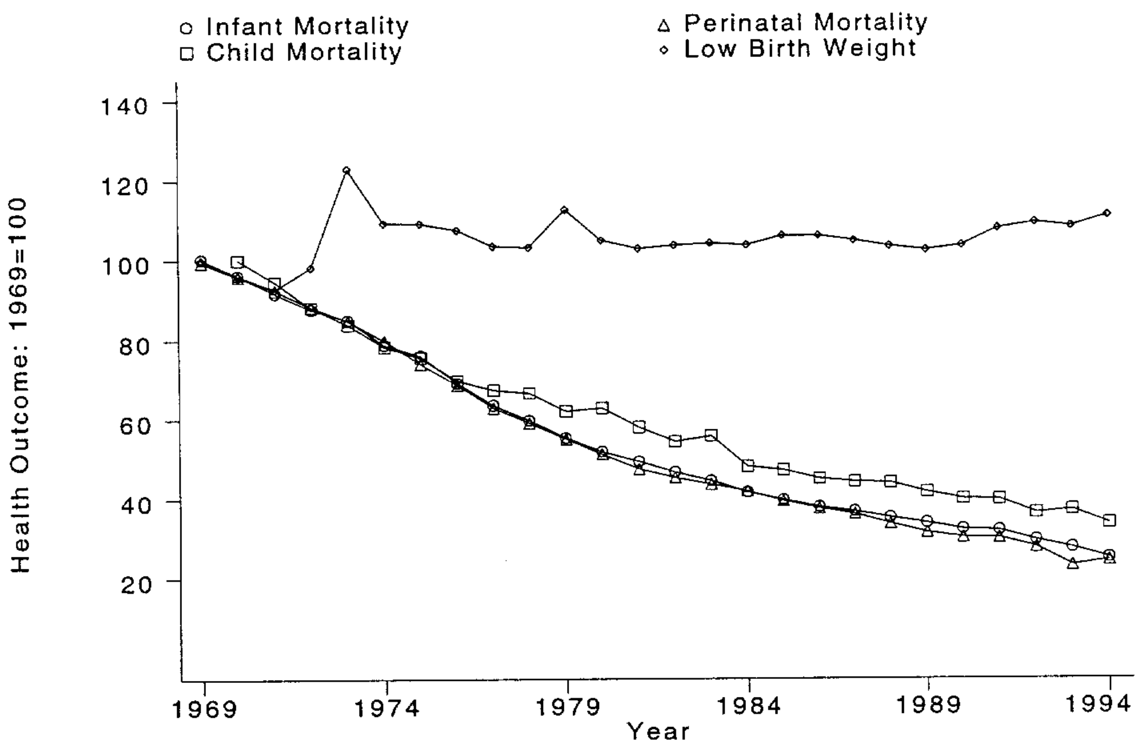

Figure 1b: Trends in Child Health Outcomes

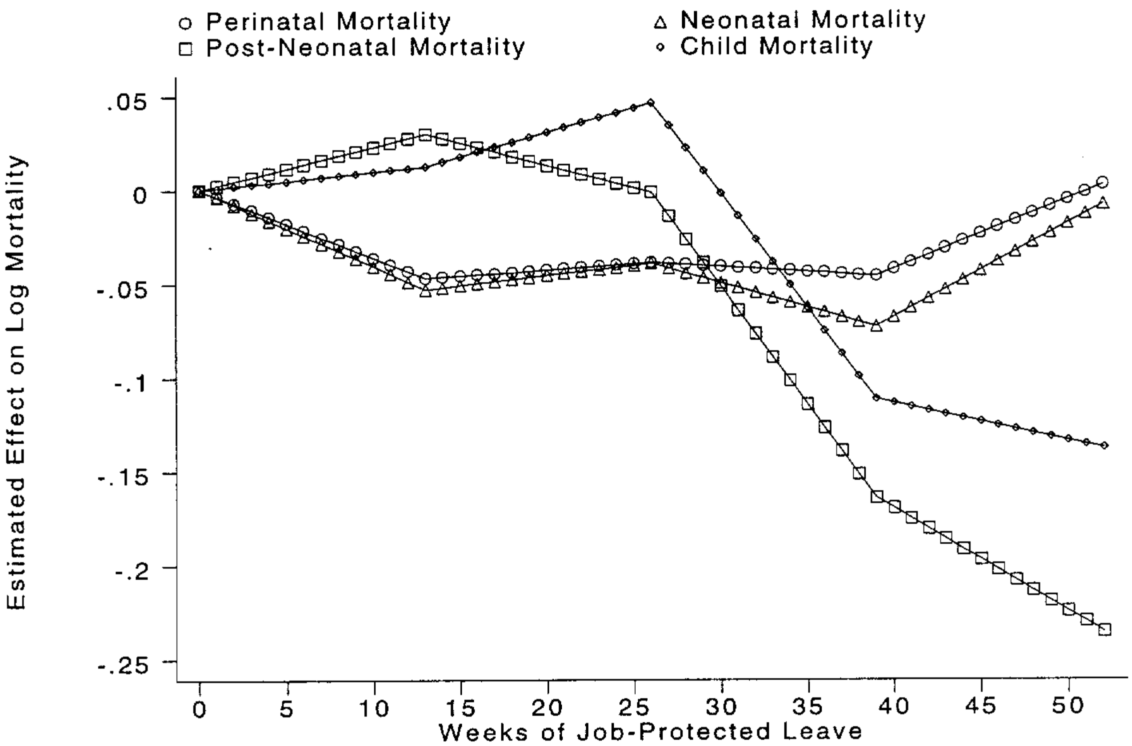

Figure 2a: Parental Leave Effects in Models Without Supplemental Regressors

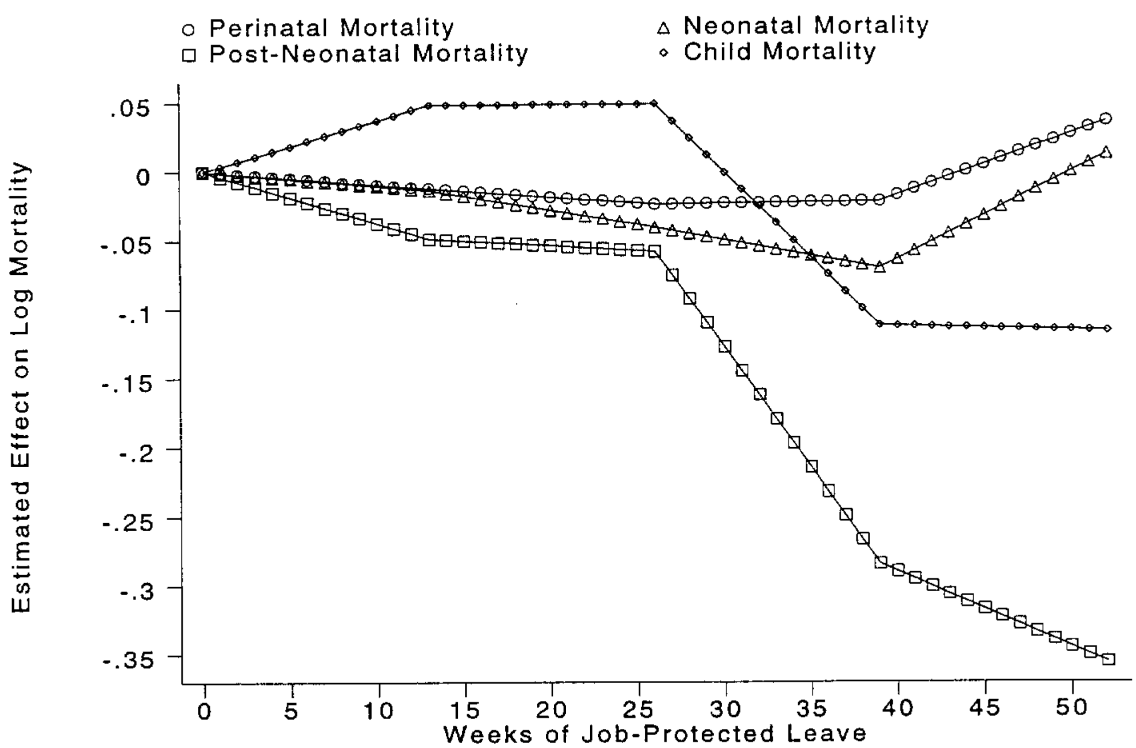

Figure 2b: Parental Leave Effects in Models With Supplemental Regressors