DIW Berlin

German Institute
for Economic Research

Discussion Papers

289

Jan Ondrich,
C. Katharina Spiess and
Qing Yang

The Effect of Maternity Leave on Women's Pay in Germany 1984-1994

Berlin, July 2002

Opinions expressed in this paper are those of the author and do not necessarily reflect views of the Institute.

DIW Berlin

German Institute
for Economic Research

Königin-Luise-Str. 5
14195 Berlin,
Germany

Phone +49-30-897 89-0
Fax +49-30-897 89-200

www.diw.de

ISSN 1619-4535

THE EFFECT OF MATERNITY LEAVE ON WOMEN'S PAY IN GERMANY 1984-1994

Jan Ondrich
Center for Policy Research
Syracuse University, Syracuse, USA

C. Katharina Spiess
German Institute for Economic Research (DIW Berlin)
Berlin, Germany
and
Technical University of Berlin
Berlin, Germany

Qing Yang
Verbind
Lexington, MA, USA

February 2001
Revised May 2002

The second author gratefully acknowledges financial support by the National German Science Foundation (Deutsche Forschungsgemeinschaft) (Grant No. Wa 547/2-1). The authors thank Richard Burkhauser, Barbara Butrica, Mark Natoli, Robert Weathers, and Gert G. Wagner for their help and suggestions. Remaining errors are, of course, our own.

Abstract

In 1986 German federal parental leave and benefit policy was expanded in several ways, extending the potential duration of leave from six to ten months and paying child-rearing benefits to all new mothers regardless of their employment status before childbirth. The potential duration has increased four times since 1986 and stood at 18 months in 1991 and three years starting in 1992. This study uses log-wage difference regressions to examine the effect of leave taken by the mother on wage growth for two 5-year periods, 1984-1989 and 1989-1994. In each of the five-year periods, taking maternity leave was found to have a significant negative effect on wage growth. Point estimates imply that each month of maternity leave reduced wage growth by 1.5 percent over five years. In addition, for the second five-year period only, mothers experienced lower wage growth if they chose to stay at home rather than return to work when the allowable leave period expired: from 1989 to 1994, a half-year out of the labor force after the end of the leave period lowered wage growth by an additional 15 percent over five years.

I. Introduction

Public policy regarding parental leave should take into account several stylized facts. Parental leave and benefit policies generally encourage the continued labor force attachment of mothers, and in the absence of such policies the demands of infant care can result in a mother's complete withdrawal from the labor market. For the employer these policies enable the retention of human capital, saving the costs of hiring and training new workers. Parental leave and benefit policies also have their costs, however. Job protection increases labor market inflexibility and benefits paid by the firm increase the cost of labor. Since mothers take the parental leave in most cases rather than fathers, these policies may result in wage discrimination against women of childbearing age.

National legislation concerning parental leave has not always been passed in countries with the largest fraction of mothers who work. For example, the employment-to-population rate for wives with children under six has been historically higher in the United States than in Germany, 53.0 percent vs. 34.3 percent in 1988,[1] and, more generally, Europe. Nonetheless, there has been a longer tradition of federal maternity leave and benefit policy in Germany and other European nations than in the United States. The first national legislation on parental leave in Germany took effect in 1979, whereas the Family and Medical Leave Act (FMLA) of 1993 represents the first U.S. legislation concerning parental leave.

Ruhm (1998) reports that for the pre-FMLA period in the United States maternity leave was available to 37 percent of women working full time in medium and large establishments (more than 100 workers) and 18 percent of those working for small employers, but that paid leave was available to only 1 and 2 percent respectively. He gives two possible explanations for these low coverage rates: market imperfections limit the provision of paid and unpaid parental leave; and most workers believe that the costs of entitlements exceed the benefits. To examine the second point Gruber and Krueger (1991), Gruber (1994), and Anderson and Meyer (1995) studied the consequences of employer mandates and determined the conditions for firms passing on the costs associated with parental leave to employees by reducing wages. They conclude that if workers place a lower value on benefit compensation than on wages, then the introduction of an employer mandate will result in a wage decline smaller than the costs of the benefits and a decline in total surplus.

Ruhm points out that this conclusion is misleading if either the costs associated with parental leave are financed out of government revenues (which is true in European nations for the most part), or when dynamic considerations are taken into account. In the second case, suppose that the introduction of parental leave raises the level of firm-specific human capital by allowing parents to retain their jobs. This elevates the marginal revenue product of labor and results in a smaller decline in the demand for labor in a dynamic analysis than in a static analysis. The wage decline is diminished and if the decline in labor demand is relatively small, the sum of producer and worker surplus will increase, implying an increase in economic efficiency. When leave costs are financed by the government, the result depends on whether the tax increase to finance the costs of leave is distortionary. If it is not, case the result will be qualitatively the same as in the dynamic analysis. Ruhm, however, considers the case of distortionary financing to be more realistic and in this case the conclusion is less clear.

Several studies (for example, Mincer and Polachek 1974; Mincer and Ofek 1982; and Corcoran, Duncan and Ponza 1983) find that time out of the labor force has a negative effect on wages. Dalto (1989) and Spalter-Roth and Hartmann (1990) find that women are out of work for less time and receive higher wages if employers voluntarily provide leave after childbirth. Ruhm points out, however, that it is not clear whether the results of these two studies are being driven by the leave entitlements themselves or by non-random selection by women into the jobs providing the entitlements. Waldfogel (1997) uses both log-wage regression and differenced log-wage regression to examine the effects of employment continuity over childbirth on women's wages for two young cohorts from the NLS-YW and NLSY data sets. She finds two reasons why employment continuity is associated with higher wages in both cohorts. First, women maintaining employment continuity start with higher wages; second, women who return to their previous employer have greater work experience and job tenure. Ruhm (1998) argues that estimation of the differenced log-wage regression does not completely eliminate biases caused by non-random selection into employment. He cites the fact that Waldfogel finds that women not taking leave, but having the option to take leave, experience greater wage growth than corresponding women without coverage. Ruhm performs cross-national analysis as an alternative. However, it is hard to completely justify such an approach if workers and firms within a country have heterogeneous characteristics.

Turning to the case of Sweden, the first important paper on work interruptions is by Gustafsson (1981), which replicates the Mincer and Polachek (1974) study using data on private-sector white-collar workers. Her two-stage least-squares results produce a significantly negative coefficient for time out of the labor force on log wages. A more recent study by Albrecht, Edin, Sundstrom, and Vroman (1999) uses the Swedish Family and Work data set, collected by Statistics Sweden starting in 1992, to estimate the depreciation effect of employment interruptions on wages. They conclude that job interruptions in the form of home time (time spent out of the labor force) and time spent unemployed have a significant negative effect on women's wages while time spent on parental leave has none. These effects are present for women only and do not occur for men. When the differenced log-wage specification is used, however, job interruptions are no longer found to have an effect for women.

Moving finally to the case of Germany, Mavromaras and Rudolph (1997) examine gender wage discrimination upon re-employment in Germany using the official micro-statistics of the German Employment Office (the IAB 1% Employment Sample). They find that although total discrimination upon re-employment declines over time, the portion directly attributable to hiring has increased. They suggest that these results are due to employers switching to discriminatory hiring practices, which are less likely to be detected and are not directly covered by existing legislation. More recently, Beblo and Wolf (2000) use the 1998 cross-section of the German Socio-Economic Panel (GSOEP) to examine the hours and participation decisions, as well as wage outcomes, of 560 women aged 30 to 55. The authors examine the role of both career interruptions and the decision to participate part-time on the wages of women. Beblo and Wolf find that a three-year break starting at age 30 leads to a wage loss of about 1.2 marks per hour.

The study by Mavromaras and Rudolph excludes from consideration employment interruptions due to parental leave, while the study by Beblo and Wolf does not distinguish parental leave from other types of employment interruption. Yet it is possible that parental leave will have a stronger effect than other types of employment interruption on the subsequent wage growth of the mother. Germany has virtually universal parental leave and benefit coverage—exclusions exist only for the self-employed and those without an existing job contract. In 1986 the German federal government expanded its maternity leave and benefit policy in several ways. It extended the potential duration of the leave from 6 to 10 months and started paying child-rearing benefits to all new mothers regardless of their employment status before childbirth. The potential duration has increased four times since 1986 and stood at 18 months in 1991 and three years in 1992.

Although the German employer is not liable for a maternity benefit except for a short interval at childbirth, the increase in the potential duration poses other problems. During the parental leave the firm must cover the position of the mother on leave with a temporary worker, whose contract, in principle, ends when the mother returns to work. If the firm cannot create a new position for the temporary worker, she must leave the firm and her accrued firm-specific human capital (up to three years since 1992) goes to waste. Moreover, since the temporary worker assumes that her job will terminate when the mother returns, it is likely that the temporary worker will leave the firm before the actual return of the mother to accept permanent employment elsewhere. This aggravates the firm's problem of keeping the position filled.

Thus, while the employer's mandate for parental benefits may not be a severe burden for a firm with an employee absent on leave, the (uncertain) costs of hiring and training temporary workers must be added to the leave costs to the firm. It is plausible that firms try to recoup these costs from the returning mother by reducing her future wage growth.

This study uses the GSOEP to examine the effect of parental leave taken on the wage growth of a German mother. This will be accomplished using differenced log-wage regressions for two periods, 1984-1989 and 1989-1994. We chose the year 1989 as the split because it conveniently divides the potential durations. Mothers in the first period have the shorter potential durations of parental leave of 6, 10, or 12 months, while mothers in the second period have longer potential durations of 15 months, 18 months or three years.

The differenced log-wage regressions include controls for differences in the levels and squares of age, education, years of labor force experience, and job tenure. The regressions also include both the woman's labor force characteristics between, but not including, the boundary years of the interval and separate variables for her characteristics in the boundary years. Job tenure measures the number of years since the woman joined the firm and therefore does not measure her accrued on-the-job firm-specific experience if she has taken leave. Accordingly, the effect of the difference between tenure and firm-specific experience is captured in the effect of the actual cumulative duration of leave between the boundary years.

This means that the effect of the actual cumulative duration of leave taken does not measure the net benefit (cost) of the available leave compared to a situation in which the leave is not available. Rather it measures the net cost of the leave compared to a situation in which the leave is not taken. Certainly the first measure is the appropriate one if one is interested in measuring the benefits of the parental leave policy, conditional on the fertility choice of individuals. However, the second measure may be of interest in a low-fertility setting in which the choices of fertility and duration of leave are both left open.

The remainder of this study is organized as follows. In section II we give an overview of the German federal parental leave and benefit policies and the changes in these policies since the early eighties. In section III, we present the differenced log-wage regression methodology. In section IV we describe the data and the variables included in the analysis. We present the results in section V and provide conclusions in section VI.

II. Parental Leave and Benefit Policies in Germany

German parental leave and benefit policy underwent sweeping changes in 1986. Before 1986 policy focused on mothers. In 1986 fathers became eligible for a leave as well; moreover, parental leave and benefit policy a more important instrument in providing child care for infants.

Maternity Leave and Benefit Policy before 1986

Employed mothers in Germany have been eligible for maternity leave and benefits since 1979. The German mother-protection law (*Mutterschutzgesetz)*, the only federal legislation in effect until 1986, contains four important regulations:

- providing employed women with protection against dismissal during pregnancy and four months after delivery;
- prohibiting work for new mothers for a period of eight weeks after childbirth—the "*Mutterschutz*" (mother protection) regulation;
- entitling mothers engaged in paid work (excluding self-employed mothers) to a protected maternity leave of four months, from the end of the mother-protection period (eight weeks after childbirth) until the child is six months old;
- entitling mothers to a maternity benefit for the six months after childbirth.

From 1979 to 1985 the benefit amount was based on average income earned in the three months of work immediately before the birth of the child. The range in the initial eight-week mother-protection period was from a minimum of DM3.50 to a maximum of DM25 per day. The mother also received a generous supplement paid by the employer during the mother-protection period. The employer paid the difference between the maternity benefit and the average income earned by the mother in the three months of work before childbirth. The mother-protection period ends two months after childbirth; from that point on, the maximum benefit was DM17 per day.

Parental Leave and Child Rearing Benefit Policy since 1986

Beginning in 1986 there were major changes in German parental leave and benefit policy. In 1985 the German Parliament passed the federal child-rearing benefit law (*Bundeserziehungsgeldgesetz*). With this law the job of one of the working parents became protected for a period of ten months after the birth of a child—eight months of protected leave for the new parent after the end of the two-month mother-protection period instead of the four months of protected leave available only to the new mother before 1986. Benefit provision also changed. A parent who stayed at home to care for the newborn child became entitled to a child-rearing benefit independent of that parent's previous employment status. The entitlement period lasted until the child was ten months old. Before 1986 only the mother could collect a benefit;

now either parent not involved in full time work could collect the benefit (full time work is defined as 20 hours per week or more). The benefit amount for each of the first six months, including the two-month mother-protection period, became DM600; the new mother continued to receive the employer supplement provided by the mother-protection law. From the seventh month on the amount of the child-rearing benefit paid for each child depended on annual net family income two years before the birth of the child; furthermore, the benefit was reduced on a sliding scale basis. For a two-parent household no benefit was available if annual net family income exceeded DM29400; the upper limit for a single-parent household was DM23700 per year. Each additional child increased the upper limit by DM4200. The child-rearing benefit is tax-free and its receipt does not require German citizenship.

The next major change in federal parental leave and benefit laws occurred in 1988. Both the benefit entitlement period and the parental leave period increased from 10 to 12 months, giving the new mother 10 months of protected leave after the mother-protection period ends. There were further increases in the total period of benefit entitlement and parental leave from 12 to 15 months in 1989, and from 15 to 18 months in 1990. The parental leave period was lengthened again in 1992—one parent at a time could now obtain parental leave with job protection until the third birthday of a child. For the first time however, the period of entitlement for the child-rearing benefit was not lengthened *in tandem* with the parental leave—entitlement for the benefit remained at 18 months.

The next change in the law increased the eligibility period for the child-rearing benefit from 18 to 24 months in 1993 without changing the potential duration of parental leave. In 1994 legislation introduced an upper limit on net annual family income for receipt of the child-rearing benefit for the first six months after the birth of the child—the legislation did not affect receipt of the supplement paid by the employer during the mother-protection period. The upper limit for receipt of the child-rearing benefit is DM100,000 for couples and DM75,000 for single parents. Furthermore, the rules for determining net annual family income changed at this time; these rules affected actual receipt of the child-rearing benefit for both the first six months after the birth of the child and the remainder of the benefit entitlement period. In 2001 the upper income limits for receipt of the child-rearing benefit were raised. A parent on parental leave could work part time (not exceeding 30 hours per week) without a change in his or her leave status and it became possible for a mother and a father to be on parental leave simultaneously (see Bertelsmann Foundation 2000).

Besides these federal regulations, during our sample period, seven German states, Baden-Wuerttemberg, Bavaria, Berlin, and Rhineland-Palatina in West Germany, and Mecklenburg-Vorpommern, Saxony and Thuringia in East Germany, have additional child-rearing benefit periods as well as recommended, but non-mandatory, extensions of the federal protection period of six to 12 months (see Rosenschon 2001). Also, some larger companies in Germany provide a parental leave longer than the federal law requires (see Bundesminister fuer Jugend, Familie, Frauen und Gesundheit [Federal Minister for Youth, Family, Women and Health] 1989).

III. Methodology

The starting point for the estimation methodology is the specification of a Mincer-type relationship between log wages and relevant human capital covariates at two time points, t and $t+k$. For a given individual, the log-wage regression can be written as

$$\ln w_j = \alpha_j + \beta' X_j + \varepsilon_j, \quad \text{for } j = t,\ t+k \ , \tag{1}$$

where X_j is the vector of covariates consisting of levels and squares of age, education, years of labor force participation, and years of job tenure, β is the coefficient vector, α_j is the cumulative effect to time j of other non-random determinants of wages inclusive of a permanent component (an individual-specific fixed effect), and ε_j is a random disturbance.

Subtracting the level of each variable at t from its level at $t+k$ and letting Δ stand for the difference operator yields

$$\Delta \ln(w) = \Delta\alpha + \beta'\Delta X + \Delta\varepsilon, \tag{2}$$

where $\Delta\alpha = \alpha_{t+k} - \alpha_t$ is parameterized as

$$\Delta\alpha = \delta + \gamma' Z. \tag{3}$$

In equation (3), δ is an intercept representing the pure effect of the time k between observations and Z is a vector of variables describing the work interruptions between time t and $t+k$. Combining equations (2) and (3) yields the estimating equation

$$\Delta \ln(w) = \delta + \gamma' Z + \beta'\Delta X + \Delta\varepsilon. \tag{4}$$

Because we are interested in the effect of work interruptions within a five-year period, it is important to account also for work interruptions in the lower boundary year. Therefore, variables describing work interruptions in the years t and $t+k$ are added to the right-hand side of equation (4). Controls for occupation, change in occupation between t and $t+k$, and state of residence are also added.

The regression results for the model in equation (4) are to be used to predict the wage ratio $s_{t,t+k} = w_{t+k} / w_t$ for the women in the sample. The regression function from equation (4) is in fact an unbiased predictor of $\ln(s_{t,t+k})$, since $\ln(w_{t+k} / w_t) = \Delta \ln(w)$. Exponentiating the regression function from equation (4) does not however yield an unbiased predictor of $s_{t,t+k}$. It will in fact systematically underpredict the expectation of $s_{t,t+k}$. To see the reason for the underprediction, note first that

$$E(\exp[\delta + \gamma' Z + \beta \Delta X + \Delta\varepsilon]) \quad = \quad \exp[\delta + \gamma' Z + \beta \Delta X]\, E(\exp[\Delta\varepsilon]) \ . \tag{5}$$

Thus, exponentiating the regression function underpredicts $s_{t,t+k}$ if it can be shown that the expectation on the right-hand side of equation (5) is greater than one. This inequality follows from the convexity of the exponential function and the fact that for any convex function g,

$$E(g(X)) \quad \geq \quad g(E(X)) \tag{6}$$

by the Jensen Inequality (see Mood, Graybill, and Boes, 1974, p.72). Thus,

$$E(\exp[\Delta\varepsilon]) \ \geq \ \exp[E(\Delta\varepsilon)] \ = \ 1 \ . \tag{7}$$

To complete the computation of an unbiased predictor of $s_{t,t+k}$, it remains to evaluate the expectation on the right-hand side of equation (5). But $E(\exp[\Delta\varepsilon])$ is simply $m_{\Delta\varepsilon}(1)$, where $m_{\Delta\varepsilon}$ is the moment-generating function for $\Delta\varepsilon$. Therefore, assuming the regression model in equation (4) is normal,[2] i.e., disturbances are independent and normally distributed variables with mean zero and variance σ^2, the expectation is given by (see Fraser, 1976, p. 234)

$$E(\exp[\Delta\varepsilon]) \ = \ m_{\Delta\varepsilon}(1) \ = \ \exp[\sigma^2 / 2] \ . \tag{8}$$

In this case, $\exp[\delta + \gamma' Z + \beta \Delta X + (\sigma^2 / 2)]$ is an unbiased predictor of the ratio w_{t+k} / w_t, and, therefore, ($\exp[\delta + \gamma' Z + \beta \Delta X + (\sigma^2 / 2)] - 1$) is an unbiased predictor of the wage growth rate, $s_{t,t+k} - 1$, between t and $t + k$. Finally, the change in the wage growth rate due to a unit increase in covariate Z_j with coefficient γ_j is $s_{t,t+k}\gamma_j$. An unbiased predictor for this change in the wage growth rate is $\gamma_j \exp[\delta + \gamma' Z + \beta \Delta X + (\sigma^2 / 2)]$. All predictors will be evaluated at estimated parameter values.

IV. Data and Variables

The data on women's wages and socio-economic variables come primarily from the English Language Public Use File of the GSOEP (see Wagner, Burkhauser, and Behringer 1993), but are augmented by variables from the German version of the GSOEP (see Wagner, Schupp, and Rendtel 1994). The GSOEP is an ongoing panel study begun in 1984, when a sample of 5,921 households was selected to represent the population of West Germany, including the five largest groups of foreigners. East German households were added in 1990, but these data are not used in this study.

Wage data come from the 1984, 1988, 1989, 1990, 1993, and 1994 waves of the GSOEP. The sample is restricted to women between the ages of 16 and 45 who are not self-employed. The first period of analysis covers the years 1984 through 1989, while the second period of analysis covers the years 1989 through 1994. For each of the two periods the woman must have valid wage information for the lower boundary year to be included in the estimation for that period. If, in either period, valid wage information for the upper boundary year is unavailable, we use the wage information from the preceding year; for the period from 1984 to 1989 period, we use the wage information for the year following the upper boundary year if it is unavailable for both the upper boundary year and the preceding year. If wage information is obtained from a year surrounding the boundary year, covariates are taken from the same year and a dummy variable on the right-hand side of the regression equation is switched on.

To ensure that the results are not being driven by observations on women with a weak attachment to the labor force, observations are excluded from the estimation if the woman reports working less than 10 hours per week in either boundary year. To minimize the effect of potential measurement error in the wage rate, the observation is excluded if the woman reports a wage rate greater than DM200 per hour in either boundary year.[3] The final samples consist of 759 women for the period from 1984 to 1989 and 769 women for the period from 1989 to 1994. In the earlier period, 680 women work for at least 6 months in each boundary year, while the number is 665 for the second period. The wage-rate mean for the lower boundary year in the earlier period is DM13.72, while for the upper boundary year it is DM17.90. The corresponding numbers for the later period are DM15.88 and DM22.29. The mean of the growth rate of wages is substantially higher than the median in both periods (58.3 compared to 31.2 percent in the earlier period and 61.5 compared to 38.1 percent in the later period).[4] The mean and standard deviation of the dependent variable are 0.331 and 0.477 for the earlier period; for the later period, the dependent variable mean and standard deviation are 0.373 and 0.433. Exact definitions of the covariates appear in Table 1, while Table 2 (Table 5) gives their means and standard deviations for the earlier (later) period.

V. Results

The differenced log-wage regression results for the period from 1984 to 1989 are presented in Table 3. Models 1 and 2 are run for all women who meet the criteria of the preceding section, while Models 3 and 4 include only those who work for at least 6 months in each boundary year. Models 1 and 3 include state fixed effects, while Models 2 and 4 do not. All specifications include additional covariates for which coefficient estimates have not been reported.

These additional covariates include one-digit ISCO occupation codes for the lower boundary year to control for the possibility that wage growth varies by occupation. The specifications also include additional covariates, none of which turn out to have a significant coefficient, giving the number of births and number of months of non-employment for each of the boundary years. Finally, each specification has an indicator for whether a missing wage rate for 1989 is filled with a value from 1990. It is not necessary to add a second dummy for whether a missing wage rate for 1989 is filled with a value from the preceding year—this new variable would be perfectly collinear with the differenced age variable.

Most of the action in the differenced right-hand side variables comes from age and education and their squares, all of which are significant at the 5 percent level based on a two-tailed test in Models 1 and 2. The combined effect of the linear and squared term for age suggests that wages increase with age for all women in our sample. Mavromaras and Rudolph (1997) and Waldfogel (1997) find qualitatively similar results for the effect of age. The combined effect of the two terms for education suggests that wages increase with up to 16 years of education. The job tenure variables are insignificant; Mavromaras and Rudolph do find significance for job tenure, but their sample consists of 12,406 women.

The years of labor force participation variables appear to have the wrong signs. The reason for this is a strong collinearity with the age variable. Years of labor force experience is the number of calendar years for which there is any labor force participation by the woman. Yet it is extremely likely that most women will work for at least a short period of time in most years. This gives labor force experience the flavor of a potential experience variable and results in the collinearity with age. This collinearity is likely to lead to imprecision in the coefficient estimates for age and labor force experience, but should not otherwise affect the predictive ability of the models.

Let us turn to the level variables, i.e., the variables that comprise Z and that are not differenced. The number of months of leave taken is hypothesized to have a negative effect on log wages. The coefficient estimate for this variable is in fact negative, and is significant at the 5 percent level based on a one-tailed test in Models 1 and 2. The results are somewhat stronger in

Models 3 and 4, in which women worked at least six months in each boundary year. The change in occupation variable is insignificant and dropping the occupation variable from the analysis has no appreciable effect on the remaining results. The remaining level variables are insignificant.

In Table 4 we calculate the marginal effect of a month of parental leave on the wage growth rate of women at 6, 10, 12, and 20 months. The remaining covariates are evaluated at their sample means. The marginal effects are consistently in the range of -1.5 percent, with slightly higher figures in Models 3 and 4. This means that a year of parental leave lowers the woman's wage growth rate from a predicted mean of 53 percent to 35 percent.

The regression results for the period from 1989 to 1994 are presented in Table 6. Models 5 and 6 include all women who meet the general criteria; Models 7 and 8, on the other hand, include only those women with at least six months of work in each boundary year. All specifications have the one-digit ISCO codes for the lower boundary year and all have variables giving the number of births and number of months of non-employment in each boundary year (a birth in the upper boundary year has a significantly negative impact on log wage differences at the 1 percent level). Only Models 5 and 7 include state fixed effects. The filling of a missing wage rate for 1994 with a value from 1993 is captured by the differenced age variable.

The significant differenced covariates are age, education, education squared, and labor force experience. Coefficient estimates for both education and education squared are significant at the 1 percent level in all four models. The coefficient estimate for age is significant at the 1 percent level in Models 5 and 6 and at the 5 percent level in Models 7 and 8. Finally, the coefficient estimate for labor force experience is significantly negative at the 5 percent level. In any case, the estimated total effects for age and education dominate the total effect of labor force experience. The job tenure variables are insignificant.

Turning to the level variables, the coefficient estimate for change in occupation is insignificant as before. Once again, in all four models the estimate of the coefficient for months of leave is significantly negative at the 1 percent level based on a one-tailed test. Moreover, the point estimates are virtually identical to those for the earlier period from 1984 to 1989. While the remaining *Z* variables all had insignificant coefficient estimates for the earlier period, in the later period from 1989 to 1994 the coefficient estimate for the number of months of post-leave non-employment is significantly negative at the 5 percent level. Moreover, depending on the model, the coefficient estimate is double or more than double the coefficient estimate for the months of leave variable. Because potential length of leave must always be longer in the later period, the presence of any post-leave non-employment implies the combination of a lost job and a significant portion of the period out of the labor force. Mothers who work in the upper boundary year in spite of this are likely to be working for substantially lower than average wages.

In Table 7 we calculate the marginal effect of a month of parental leave on the wage growth rate of women at 6, 10, 12, and 18 months of leave. We also calculate the marginal effect of a month of post-leave non-employment evaluated at the first and sixth months. The remaining covariates are evaluated at their sample means. The marginal effects for months of parental leave are uniformly in the range of -1.3 to -1.5. This means that a year of parental leave will lower the wage growth rate from a predicted mean of 56 percent to 38 percent. Eighteen months of leave will virtually cut the wage growth rate in half, from 56 to 29 percent.

The marginal effect of a month of post-leave non-employment is somewhat larger than the marginal effect for a month of leave. The most conservative estimates here are -2.3 percent for the first month of post-leave non-employment (based on a potential duration of 18 months) and -2.1 percent for the sixth month of post-leave non-employment. This means that 18 months of leave followed by six months of post-leave non-employment will cut the wage growth rate by more than two-thirds, from 56 percent to 16 percent.

While it may be premature to discount the effect of job tenure and years of labor force participation on the wage rates of women in Germany over the period of time from 1984 to 1994, the effect of parental leave seems robust to a wide variety of specifications in both halves of the interval. An extra month of parental leave appears to decrease the growth rate of wages by 1.5 percentage points for a five-year interval. In the absence of affordable and generally available day care (see Kreyenfeld et al. 2001), a working prospective mother with no option but to take the full extent of leave, should she have a child, must make a difficult decision concerning career versus family.

There are several points to consider before it can be concluded that taking a longer parental leave has a significant long-run effect on wage growth, and more research is needed. First, it is entirely possible that over a period greater than five years, the effect of maternity leave on wage growth diminishes. Second, the use of ordinary least squares to estimate the determinants of the log wage difference deserves closer scrutiny. The mean value of the log wage difference can be close to 90 percent higher than the median value of the log wage difference, suggesting a role for median and quantile regression in evaluating the potential influence of outliers on the parameter estimates.

An interesting application of median wage regressions in the labor economics is found in Johnson, Kitamura, and Neal (2000). An interesting aspect of the study by Johnson et al. is the suggestion that replacing missing values of the truncated dependent variable with a value of zero will, under certain conditions, produce consistent estimates of the wage equation parameters. This brings us to the final consideration—the effect of selectivity on the estimates of the effect of an extra month of maternity leave on wage rates. The sample that we use is a selected sample because only women who participate in the labor force in both boundary years of an interval are included the log-wage regressions.

The first issue in dealing with the effect of selectivity on the estimated coefficients of the log-wage regression is to recognize that the log-wage regression function (conditional expectation function or CEF) for a selected subpopulation is nonlinear in parameters even if the log-wage regression function for entire population is linear in parameters. Poirier and Melino (1978) showed that the slopes of the nonlinear CEF for the selected subpopulation, evaluated at any value of the vector of regressors, are proportionally attenuated (proportionally biased asymptotically toward zero in absolute value) with respect to the slopes of the nonlinear CEF for the entire population.

Another question that might be asked is: What is the effect of selection on the Ordinary Least Squares (OLS) coefficient estimates of the linear-in-parameters population log-wage regression specification (linear regression function or LRF slopes)? Cain and Watts (1973, pp.342-343) and Hausman and Wise (1977, p.935) suggest and provide support for the conjecture that the LRF slopes will also be proportionally attenuated. However, Goldberger (1981) demonstrates that a sufficient condition for the proportional attenuation of LRF slopes is that the dependent variable and explanatory variables in the regression are multinormally distributed. Moreover, Goldberger presents an example where the dependent variable and explanatory variables are not jointly normal, with the result that the LRF slopes are not attenuated.

If in fact the slope coefficient for the number of months of maternity leave in the log-wage regression is attenuated, the estimate of a 1.5 percent decline in wages for each month of leave understates the true value. We now present an argument suggesting that attenuation is a reasonable outcome. Because affordable day care for infants and toddlers is very scarce in West

Germany, mothers of young children often find it difficult to combine work and raising children. In 1994, there were only two day-care slots per 100 children under the age of three. Four years later, this number had increased by 50 percent, to three day-care slots per 100 children under the age of three. Therefore, because of this scarcity of day care, West German mothers usually leave the labor force after the birth of their first child (see Kreyenfeld, et al. 2001, and Spiess and Buechel 2002). As a result the labor force participation rate of mothers with young children is substantially lower than the corresponding rates for non-mothers and mothers with only older children (see Holst and Schupp 2001, Table 3).

The relevance of these facts for an analysis of selectivity bias can be seen in Figure 1. The argument is simplified by assuming three stylized groups of women. The women in Group 1 either have no children, and therefore zero months of leave, or have had a child, but took only an extremely short leave. Women who had a child and took a long leave (all leaves in this example are either long or extremely short) are either in Group 2 or Group 3. The mothers in Group 2 have high wages, while the mothers in Group 3 have low wages. Mothers whose wage growth does not exceed a certain reservation level cannot afford day care and no longer participate in the labor market. (The reservation wage growth rate is drawn to be constant in Figure 1 but the argument remains valid if the reservation growth rate varies between the high wage changes of mothers in Group 2 and the low wage changes of mothers in Group 3). The line labeled as the true line assumes that there is no selection from the population and the line labeled as the estimated line assumes that the Group 3 mothers do not provide information to the LRF. The population LRF slope is the slope of the true line and the LRF slope for the selected subpopulation is the slope of the estimated line. It is clear that the selectivity attenuates the LRF slope for months of leave. Our conclusion is that the lack of affordable day care in West

Germany is likely to result in a downward bias for the effect of each month of maternity leave on wage growth.

VI. Conclusions

This study examines the effect of parental leave taken by German mothers on their wage growth for the period from 1984 to 1994 using differenced log-wage regression to remove the individual-specific permanent component for log wages. The ten-year period was divided into two five-year periods to examine whether any negative effect grew as the German federal leave and benefit policy became more liberal. Much of the effect of the standard first-differenced covariates was captured by age and education alone. Nevertheless, over a wide variety of specifications, the estimated marginal effect on the wage growth rate of an extra month of parental leave was -1.5 percent for both of the five-year periods. The estimated marginal effect implies that a year of parental leave cuts the five-year wage growth rate by one-third and 18 months of leave cuts the growth rate by one-half. In the later five-year period from 1989 to 1994 but not in the earlier period from 1984 to 1989, the length of post-leave non-employment was found to have a significantly negative impact on wage growth as well. This may be an artifact of the increased potential duration of maternity leave in the later period—the presence of any post-leave non-employment would imply non-participation in the labor force over most of the interval. It is possible that this effect would disappear if the time period under consideration were extended.

The use of ordinary least squares to estimate the determinants of log wage differences should be more closely examined. In the estimation sample the mean value of the log-wage difference was close to 90 percent higher than the median, suggesting a role for median and quantile regression in evaluating the influence of outliers on the parameter estimates. We examined the effect of selectivity bias on our conclusions and believe that correcting for selectivity bias would increase our estimate of the penalty that mothers face for each month of maternity leave that they take.

The empirical work in this study has significant policy implications. One important goal of parental or maternity leave policy is to make it easier for women to combine work with starting a family. Federal parental leave policy in Germany has focused on providing job protection to new mothers for longer periods of time. The longer periods of potential leave mean that, initially, that the new mother is insulated from the general scarcity of affordable day care in Germany. But eventually, the new mother who wants to work must return to work. The situation she faces at work on her return may be a critical factor in her decision to start a family. The empirical work in this study suggests that a working woman who has no option but to take the full parental leave should she have a child, must make a difficult decision concerning career versus family.

Endnotes

1. The United States figure comes from the United States Department of Commerce, Bureau of the Census (1990) and the German figure comes from Statistisches Bundesamt (1995).

2. Duan (1983) and Wooldridge (1999) discuss methods for evaluating the expectation when the distribution of disturbances is potentially non-normal. Duan suggests using an estimation method called smearing, while Wooldridge presents an algorithm using non-linear least squares.

3. Deleting observations from the bottom decile of wages in each boundary year did not alter the main results.

4. Consumer prices increased 6.16 percent from 1984 to 1989 and 18.14 percent from 1989 to 1994.

References

Albrecht, James W., Per-Anders Edin, Marianne Sundstrom, and Susan B. Vroman. 1999. "Career Interruptions and Subsequent Earnings." *Journal of Human Resources*, 34:294-311.

Anderson, Patricia M. and Bruce Meyer. 1995. "The Incidence of a Firm-Varying Payroll Tax: The Case of Unemployment Insurance," NBER Working Paper #5201. Cambridge, MA: National Bureau of Economic Research.

Beblo, Miriam and Elke Wolf. 2000. "How Much Does a Year Off Cost? Estimating the Wage Effects of Employment Breaks and Part-Time Periods," ZEW Discussion Paper No. 00-69. Mannheim: Centre for European Economic Research.

Bertelsmann Foundation. 2000. *International Reformmonitor, Issue* 3, Gütersloh, Germany: Bertelsmann.

Bundesminister fuer Jugend, Familie, Frauen und Gesundheit [Minister for Youth, Family,Woman and Health] (ed.). 1989. "*Erziehungsgeld, Erziehungsurlaub und Anrechnung von Erziehungszeiten in der Rentenversicherung* [Child-Rearing Benefit, Parental Leave and Credits for Child Rearing in the Social Security System]," Schriftenreihe des Bundesministers fuer Jugend, Familie, Frauen und Gesundheit, Band 243. Berlin: Kohlhammer.

Cain, Glen G. and Harold Watts. 1973. "Toward a Summary and Synthesis of the Evidence." Chapter 9 in *Income Maintenance and Labor Supply: Econometric Studies*, Glen G. Cain and Harold W. Watts, eds., 328-367. Chicago: Markham.

Corcoran, Mary, Greg Duncan, and M. Ponza. 1983. "A Longitudinal Analysis of White Women's Wages," *Journal of Human Resources*, 18:497-520.

Dalto, Guy C. 1989. "A Structural Approach to Women's Hometime and Experience-Earnings Profiles: Maternity Leave and Public Policy," *Population Research and Policy Review*, 8:247-266.

Duan, Naihua. 1983. "Smearing Estimate: A Nonparametric Retransformation Method," *Journal of the American Statistical Association, 78*:605-610.

Fraser, D.A.S. 1976. *Probability and Statistics: Theory and Applications*. North Scituate, MA: Duxbury.

Goldberger, Arthur S. 1981. "Linear Regression after Selection," *Journal of Econometrics*, 15:357-366.

Gruber, Jonathan. 1994. "The Incidence of Mandated Maternity Benefits," *American Economic Review*, 84:622-641.

Gruber, Jonathan and Alan B. Krueger. 1991. "The Incidence of Mandated Employer-Provided Health Insurance." In *Tax Policy and the Economy*, David Bradford, ed., 111-143. Cambridge, MA: National Bureau of Economic Research.

Gustafsson, Siv. 1981. "Male-Female Lifetime Earnings Differentials." In *Studies in Labor Market Behavior: Sweden and the United States*, Gunnar Eliasson, Bertil Holmlund and Frank Stafford, eds., 235-268. Stockholm: Industrial Institute for Social and Economic Research.

Hausman, Jerry A. and David A. Wise. 1977. "Social Experimentation, Truncated Distributions, and Efficient Estimation," *Econometrica*, 45:919-928.

Holst, Elke and Juergen Schupp. 2001. "Erwerbsverhalten von Frauen: Trotz Annaeherung immer noch deutliche Unterschiede zwischen Ost und West." In *DIW-Wochenbericht* 42/2001. Berlin: DIW.

Johnson, William, Yuichi Kitamura, and Derek Neal. 2000. "Evaluating a Simple Method for Estimating Black-White Gaps in Median Wages," *American Economic Review*, 90:339-43.

Kreyenfeld, Michaela, C. Katharina Spiess, and Gert G. Wagner. 2001. *Finanzierungs- und Organisationsmodelle institutioneller Kinderbetreuung: Analysen zum Status quound Vorschlaege zur Reform*. Neuwied, Germany: Luchterhand.

Mavromaras, Kostas G., and Helmut Rudolph. 1997. "Wage Discrimination in the Reemployment Process." *Journal of Human Resources*, 32:812-860.

Mincer, Jacob. 1974. *Schooling, Experience, and Earnings*. New York: National Bureau of Economic Research.

Mincer, Jacob and H. Ofek. 1982. "Interrupted Work Careers: Depreciation and Restoration of Human Capital," *Journal of Human Resources*, 17:3-24.

Mincer, Jacob and S.W. Polachek. 1974. "Family Investments in Human Capital: Earnings of Women," *Journal of Political Economy*, 82:576-608.

Mood, Alexander M., Franklin A. Graybill, and Duane C. Boes. 1974. *Introduction to the Theory of Statistics, Third Edition*. New York: McGraw-Hill.

Poirier, Dale J. and Angelo Melino. 1978. "A Note on the Interpretation of Regression Coefficients within a Class of Truncated Distributions," *Econometrica*, 46:1207-1209.

Rosenschon, Astrid. 2001. "Familienförderung in Deutschland – eine Bestandsaufnahme," Kieler Arbeitspapier Nr. 1071, University of Kiel.

Ruhm, Christopher J. 1998. "The Economic Consequences of Parental Leave Mandates: Lessons from Europe." *Quarterly Journal of Economics*, 113:285-317.

Spalter-Roth, Roberta and Heidi Hartmann. 1990. *Unnecessary Losses: Costs to Americans of the Lack of Family and Medical Leave*. Washington, DC: Institute for Women's Policy Research.

Spiess, C. Katharina and Felix Buechel. 2002. "Effekte der regionalen Kindergarteninfrastruktur auf das Arbeitsangebot von Muettern." In *Wechselwirkungen zwischen Arbeitsmarkt und sozialer Sicherung* II, Winfried Schmaehl, ed. Berlin: Duncker and Humblot.

Statistisches Bundesamt. 1995. *Bevoelkerung und Erwerbstaetigkeit. Fachserie 1*, Reihe 3. Stuttgart: Metzler-Poeschel.

U.S. Department of Commerce, Bureau of the Census. 1990. *Statistical Abstract of the United States: 1990* (110th Edition). Washington, DC: Government Printing Office.

Wagner, Gert G., Richard V. Burkhauser, and Friederike Behringer. 1993. "The Syracuse University English Language Public User File of the German Socio-Economic Panel," *Journal of Human Resources*, 28: 429-433.

Wagner, Gert G., Juergen Schupp, and Ulrich Rendtel. "Das Sozio-oekonomische Panel—Methoden der Datenproduktion und -aufbereitung im Laengsschnitt." In *Mikroanalytische Grundlagen der Gesellschaftspolitik*, *Band 2—Erhebungsverfahren, Analysemethoden und Mikrosimulation,* Hauser, Richard, Notburga Ott, and Gert G. Wagner, eds., 70-111. Berlin: Akademie Verlag.

Waldfogel, Jane. 1997. "Working Mothers Then and Now: A Cross-Cohort Analysis of the Effects of Maternity Leave on Women's Pay." In *Gender and Family Issues in the Workplace*, Blau, Francine D. and Ronald G. Ehrenberg, eds., 92-126. New York: Russell Sage Foundation.

Walker, James R. 1995. "Parental Benefits and Employment and Fertility Dynamics." University of Wisconsin-Madison, and National Bureau of Economic Research. Cambridge, MA: National Bureau of Economic Research.

Wenk, D. and P. Garrett. 1992. "Having a Baby: Some Predictions of Maternal Employment Around Childbirth," *Gender and Society*, 6:49-65.

Wooldridge, Jeffrey M. 1999. Introductory Econometrics: A Modern Approach. Cincinnati: South-Western.

Table 1. Variable Definitions

Variable	Definition
Age	Age in years. Entered in log wage difference estimation as difference between ending year and starting year.
Age Squared	Age in years squared. Entered in log wage difference estimation as difference between ending year and starting year.
Education	Education in years. Entered in log wage difference estimation as difference between ending year and starting year.
Education Squared	Education in years squared. Entered in log wage difference estimation as difference between ending year and starting year.
Labor Force Experience	Labor force experience in years. Entered in log wage difference estimation as difference between ending year and starting year.
Labor Force Experience Squared	Labor force experience in years squared. Entered in log wage difference estimation as difference between ending year and starting year.
Job Tenure	Years with firm. Entered in log wage difference estimation as difference between ending year and starting year.
Job Tenure Squared	Years with firm squared. Entered in log wage difference estimation as difference between ending year and starting year.
Change in Occupation	Indicator variable for change in occupation (as given by International Standard Classification of Occupation (ISCO) code) between starting year and ending year.
Months of Leave	Months of maternity leave between starting year and ending year (not including months in starting or ending years).
Post-Leave Non-Employment Months	Months of non-employment following maternity leave between starting year and ending year (not including months in starting or ending years).
Other Non-Employment Months	Months of non-employment not following maternity leave between starting year and ending year (not including months in starting or ending years).
Births with Minimal Leave	Number of births between (but not including) starting and ending years in which mother used less than two months leave.

Table 2. Variable Means and Standard Deviations: 1984 to 1989 Period
(standard deviation in parenthesis)

	1984	1989	Difference
Differenced Variable			
Age	31.033	35.883	4.850
	(8.271)	(8.298)	(0.358)
Age Squared	1,031.36	1,356.34	324.974
	(520.390)	(601.768)	(85.303)
Education	11.503	11.660	0.157
	(2.212)	(2.240)	(0.657)
Education Squared	137.215	140.962	3.747
	(61.407)	(62.747)	(17.626)
Labor Force Experience	12.640	17.431	4.791
	(7.309)	(7.374)	(0.516)
Labor Force Experience Squared	213.130	358.132	145.001
	(211.878)	(282.322)	(73.950)
Job Tenure	6.277	9.662	3.386
	(5.646)	(6.540)	(3.209)
Job Tenure Squared	71.231	136.076	64.845
	(120.273)	(173.980)	(73.925)
Level Variable			
Change in Occupation		0.209	
		(0.407)	
Months of Leave		0.933	
		(3.085)	
Post-Leave Non-Employment Months		0.476	
		(3.226)	
Other Non-Employment Months		1.560	
		(5.600)	
Births with Minimal Leave		0.022	
		(0.148)	

Table 3. Log Wage Difference Regressions: 1984 to 1989 Period
(t-statistic in parenthesis)

	Model 1[a]	Model 2	Model 3[a]	Model 4
Differenced Variable				
Age	0.192	0.200	0.149	0.160
	(2.146)	(2.267)	(1.560)	(1.695)
Age Squared	-0.002	-0.002	-0.001	-0.001
	(-4.336)	(-4.403)	(-3.737)	(-3.541)
Education	0.811	0.807	0.737	0.743
	(6.073)	(6.147)	(4.724)	(4.840)
Education Squared	-0.026	-0.026	-0.024	-0.024
	(-5.253)	(-5.339)	(-4.324)	(-4.448)
Labor Force Experience	-0.082	-0.086	-0.063	-0.072
	(-1.135)	(-1.196)	(-0.791)	(-0.913)
Labor Force Experience Squared	0.0002	0.0002	0.0003	0.0003
	(0.575)	(0.539)	(0.725)	(0.720)
Job Tenure	0.012	0.011	0.008	0.007
	(1.420)	(1.324)	(1.060)	(0.954)
Job Tenure Squared	-0.0004	-0.0003	-0.0003	-0.0002
	(-0.987)	(-0.916)	(-0.759)	(-0.693)
Level Variable				
Intercept	0.263	0.222	0.202	0.158
	(1.089)	(0.935)	(0.794)	(0.634)
Change in Occupation	0.022	0.027	0.031	0.037
	(0.491)	(0.611)	(0.657)	(0.814)
Months of Leave	-0.011	-0.010	-0.012	-0.012
	(-1.823)	(-1.786)	(-2.004)	(-1.968)
Post-Leave Non-Employment	0.002	0.002	0.004	0.003
Months	(0.283)	(0.200)	(0.465)	(0.390)
Other Non-Employment Months	-0.004	-0.004	-0.006	-0.006
	(-0.923)	(-0.901)	(-1.146)	(-1.141)
Births with Minimal Leave	-0.028	-0.038	0.084	0.079
	(-0.254)	(-0.351)	(0.761)	(0.717)
R^2	0.197	0.192	0.143	0.138
Adjusted R^2	0.160	0.166	0.099	0.107

[a]Model includes state fixed effects.

Table 4. Effects of Marginal Month on Wage Growth Rate: 1984 to 1989 Period

Month	Model 1	Model 2	Model 3	Model 4
6	-0.015	-0.015	-0.017	-0.016
10	-0.015	-0.014	-0.016	-0.015
12	-0.014	-0.014	-0.015	-0.015
20	-0.013	-0.013	-0.014	-0.014

Table 5. Variable Means and Standard Deviations: 1989 to 1994 Period
(standard deviation in parenthesis)

	1989	1994	Difference
Differenced Variable			
Age	30.905	35.745	4.840
	(8.210)	(8.225)	(0.367)
Age Squared	1,022.43	1,345.28	322.848
	(515.913)	(596.397)	(84.442)
Education	11.618	11.791	0.173
	(2.162)	(2.150)	(0.644)
Education Squared	139.657	143.644	3.987
	(59.760)	(59.950)	(16.113)
Labor Force Experience	12.644	17.391	4.748
	(7.502)	(7.582)	(0.532)
Labor Force Experience Squared	216.074	359.873	143.798
	(218.678)	(291.317)	(75.540)
Job Tenure Squared	6.476	9.531	3.055
	(5.987)	(7.170)	(3.497)
Job Tenure	77.734	142.171	64.437
	(121.553)	(186.192)	(80.342)
Level Variable			
Change in Occupation		0.203	
		(0.402)	
Months of Leave		1.336	
		(4.673)	
Post-Leave Non-Employment Months		0.265	
		(1.988)	
Other Non-Employment Months		0.809	
		(3.472)	
Births with Minimal Leave		0.012	
		(0.119)	

Table 6. Log Wage Difference Regressions: 1989 to 1994 Period
(t-statistic in parenthesis)

	Model 5[a]	Model 6	Model 7[a]	Model 8
Differenced Variable				
Age	0.252	0.254	0.188	0.198
	(3.177)	(3.231)	(2.345)	(2.493)
Age Squared	-0.001	-0.001	-0.0002	-0.0003
	(-1.771)	(-1.841)	(-0.713)	(-0.783)
Education	0.871	0.884	0.428	0.440
	(6.687)	(6.797)	(2.747)	(2.835)
Education Squared	-0.027	-0.028	-0.014	-0.015
	(-5.241)	(-5.340)	(-2.531)	(-2.616)
Labor Force Experience	-0.147	-0.145	-0.154	-0.161
	(-2.241)	(-2.210)	(-2.265)	(-2.392)
Labor Force Experience Squared	-0.0006	-0.0006	-0.0005	-0.0004
	(-1.373)	(-1.330)	(-1.277)	(-1.141)
Job Tenure	-0.007	-0.008	-0.007	-0.008
	(-1.091)	(-1.192)	(-1.173)	(-1.130)
Job Tenure Squared	0.000	0.000	-0.000	-0.000
	(0.173)	(0.100)	(-0.164)	(-0.234)
Level Variable				
Intercept	0.269	0.282	0.426	0.407
	(0.216)	(1.322)	(2.124)	(2.064)
Change in Occupation	-0.024	-0.023	-0.013	-0.015
	(-0.597)	(-0.562)	(-0.343)	(-0.408)
Months of Leave	-0.010	-0.010	-0.011	-0.011
	(-2.440)	(-2.324)	(-2.655)	(-2.653)
Post-Leave Non-Employment Months	-0.026	-0.026	-0.020	-0.021
	(-2.492)	(-2.499)	(-2.066)	(-2.170)
Other Non-Employment Months	-0.004	-0.005	-0.013	-0.013
	(-0.882)	(-1.010)	(-2.667)	(-2.822)
Births with Minimal Leave	-0.003	-0.013	-0.013	-0.006
	(-0.025)	(-0.108)	(-0.123)	(-0.061)
R^2	0.272	0.256	0.096	0.082
Adjusted R^2	0.239	0.232	0.049	0.047

[a]Model includes state fixed effects.

Table 7. Effect of Marginal Month on Wage Growth Rate: 1989 to 1994 Period

	Model 5	Model 6	Model 7	Model 8
Leave Month				
6	-0.015	-0.014	-0.015	-0.015
10	-0.014	-0.014	-0.014	-0.014
12	-0.014	-0.013	-0.014	-0.014
18	-0.013	-0.013	-0.013	-0.013
Post-Leave Non-Employment Month[a]				
1	-0.034	-0.034	-0.023	-0.024
6	-0.030	-0.030	-0.021	-0.022

[a]Assuming leave of 18 months.

FIGURE 1: ANALYSIS OF BIAS