Parental Employment and Child Cognitive Development

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Abstract

This study investigates how parental employment affects child cognitive development. The results indicate that maternal labor supply during the first three years of the child's life has a small negative effect on the predicted verbal ability of 3 and 4 year olds and a larger detrimental impact on the reading and mathematics achievement of 5 and 6 year olds – working an extra 20 hours per week is anticipated to reduce test scores from the median to the 49th, 46th, and 47th percentiles respectively. Job-holding during the second and third years has particularly deleterious consequences if the mother works long hours or was also employed in the first year. The findings are robust to the inclusion of controls for day care arrangements or paternal employment. There is some indication that early work may be particularly costly for children in "traditional" two-parent families and the data also hint at the importance of time investments by fathers. There are two main reasons why this study provides a more negative assessment of the impact of early employment than most prior research. First, previous analyses often control relatively crudely for differences in child and household characteristics that are correlated with parental labor supply. Second, the deleterious effects are more pronounced for the reading and mathematics performance of 5 and 6 year old children than for the verbal test scores of 3 and 4 year olds that have more frequently been examined.

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The first three years of life are a particularly important period for children (Carnegie Task Force on Meeting the Needs of Young Children, 1994; Council of Economic Advisers, 1997). Recent research emphasizes the effects of early influences on brain development (Shore, 1997), and the first years are likely to be significant for the development of learning skills, self-esteem and emotional security.[1] Investments by parents during early childhood may therefore promote healthy development.

Between 1980 and 1998, the labor force participation rate of mothers with children under the age of 6 grew from 47 to 65 percent (U.S. Department of Commerce, 1999). The increase was even larger for women with infants, rising from 38 to 59 percent (Bachu and O'Connell, 2000). These changes, particularly when combined with the growth in single-parent households, suggest that adults have less time and energy to invest their offspring, with potentially deleterious effects.[2] However, increased employment may also yield benefits, most obviously by providing extra income.

This paper analyzes the relationship between parental employment and three high quality assessments of the cognitive skills of young children using data from multiple years of the National Longitudinal Survey of Youth (NLSY). The results imply a negative effect of early maternal employment. Working during the first year is associated with reductions in the verbal ability of 3 and 4 year olds that are partially (but not completely) offset by increases from job-holding during the next two years. Larger deleterious impacts are observed for the reading and mathematics achievement of 5 and 6 year olds. Labor supply in the child's second and third year is also predicted to have more harmful (or less favorable) consequences when the mother works

[1] Heckman (2000) stresses the importance of early human capital investments, stressing dynamic complementarities, whereby initial skill development fosters subsequent learning. However, the mechanisms for these effects are poorly understood and the relationship between early brain development and future outcomes remains controversial (Bruer, 1999).

[2] The proportion of children in two-parent households declined from 77 percent in 1980 to 68 percent in 1998 (U.S. Department of Commerce, 1999). Increased female employment has not been offset by substantial reductions in male work hours but fertility rates and time spent in housework have declined substantially since the 1960s (Juster & Stafford, 1991; Mayer, 1997). The time parents have available for children declined 22 hours per week (or 14 percent) between 1969 and 1999 (Council of Economic Advisers, 1999). However, Sandberg and Hofferth (2001) argue that behavioral changes by parents have prevented there from being any decrease in time actually devoted to children since the early 1980s.

long hours or was also employed in the first year. Finally, the data hint that paternal and maternal employment may affect child cognitive development in similar ways, suggesting that investments by fathers are also important.

These findings contrast with previous research predicting positive or neutral effects of parental employment during the first three years and recent studies identifying deleterious impacts of working during infancy but not thereafter. There are two main reasons for the more negative results. First, the data reveal more deleterious consequences on the reading or mathematics achievement of 5 and 6 year olds than for the verbal ability of 3 and 4 year olds that has more commonly been examined. Second, children with working parents come from relatively advantaged family backgrounds and possess attributes that promote cognitive development. Most prior analyses inadequately control for this heterogeneity and therefore substantially understate the costs of early employment. This study does a better job of accounting for the disparities and so provides a more pessimistic but more accurate assessment. However, to the extent that all confounding factors are still not held constant, these estimates may continue to understate the negative impact of parental employment.

A. Previous Research

The relationship between early maternal employment and child cognitive development has been widely studied. As detailed in Table 1, a few investigations find positive effects (e.g. Vandell & Ramanan, 1992; Parcel & Menaghan, 1994), others negative impacts (Leibowitz, 1977; Stafford, 1987; Mott 1991), and many obtain results that differ depending on the timing of work or the specific group analyzed (e.g. Desai et al., 1989; Baydar & Brooks-Gunn, 1991; Blau & Grossberg, 1992; Parcel & Menaghan, 1994).[3]

A common conclusion of the most recent (and generally carefully conducted) analyses is that maternal employment during the child's first year has a deleterious impact (e.g. James-

[3] Related research examining whether early child care harms infant-mother attachment relationships (e.g. Belsky & Rovine, 1988; Clarke-Stewart, 1989; Lamb & Sternberg, 1990; NICHD Early Child Care Research Network, 1997b) suggests that maternal employment, by increasing the use of day care, could reduce the security of attachments in some situations. However, the size and pattern of the effect is uncertain and none of the studies use nationally representative samples or contain sufficient controls for parental abilities or attitudes.

Burdumy, 1999; Han et al., 2001; Hill et al., 2001; Brooks-Gunn et al., forthcoming; Waldfogel et al., forthcoming) but often with no effect or offsetting benefits for working during the second and third year. Harvey (1999) finds that the negative impact of first year employment is temporary, whereas Han et al. (2001) and Waldfogel et al. (forthcoming) indicate greater persistence. The patterns may also vary with factors such as race, sex of the child, and family income in ways that are poorly understood. Finally, the limited research on paternal employment obtains inconclusive results (Parcel & Menaghan, 1994; Harvey, 1999; Ermisch and Francesconi, 2001; Waldfogel et al., forthcoming).

Even these inferences should be viewed as tentative because mothers holding jobs or working long hours may differ from those who do not in ways that have been inadequately accounted for. Women with observable characteristics associated with high ability, such as education levels and AFQT scores, tend to have relatively elevated employment rates (e.g. see Vandell & Ramanan, 1992; Hill et al., 2001; Waldfogel et al., forthcoming). If these advantages extend to productivity in home activities, maternal employment will be positively associated with child outcomes even absent a causal impact, suggesting that previous research may present an overly optimistic assessment of its effects.[4]

Without an exogenous source of variation in labor supply, it is difficult to completely eliminate the potential for omitted variables bias. The primary approach below is to include an usually comprehensive set of explanatory variables and to examine how the estimates change when sequentially controlling for an increasing portion of the heterogeneity. The results suggest strong complementarity between market and home abilities, implying that prior investigations have probably underestimated the costs of job holding during the child's early years.

[4] The bias could be in the opposite direction if women rapidly returning to work have less interest or ability in home production; the empirical evidence provides no support for this possibility. There are similar difficulties interpreting results of the related literature on day care (e.g. Clarke-Stewart, 1991; Field, 1991; Caughy et al., 1994). NICHD Early Child Research Network (1997a) shows that the age at which children are placed in non-parental care varies significantly with family characteristics. Those beginning care prior to 3 months of age come from families with relatively low non-maternal income whereas those placed in care at 3-5 months have the highest family income and maternal earnings. A few studies use quasi-experimental designs to control for omitted variables (e.g., Currie & Thomas, 1995). Karoly et al. (1998) provide an in-depth review of research examining early intervention programs.

B. Conceptual Framework

Economic models portray households as productive entities where parents allocate resources to maximize an objective function that includes the health and development of children as one argument. Holding marital status and household size constant more time, energy, and income are likely to be devoted to children as more of each are available to the parents.[5] Parental employment may therefore improve child outcomes by increasing incomes. However, under general conditions, market employment is also likely to lead to at least some decrease in child-related investments in time or energy. Time-diary data confirm that working mothers spend less time with children than their nonemployed counterparts (e.g. Bryan and Zick, 1996; Zick and Bryan, 1996; Bianchi, 2000; Gershuny, 2000; Hofferth, 2001; Sandberg and Hofferth, 2001), although there is considerable uncertainty about the size of this effect.[6] Long hours might also cause parents to be tired or stressed, reducing the quality of the time with children.[7]

The psychological and sociological literatures emphasize complementary pathways through which market work by parents may negatively affect child development. Belsky (1988) argues that a mother's absence during the first year of life could disrupt mother-child attachments and deprive the child of the stimulation that promotes cognitive development. Hoffman (1980) states that the stress of maternal employment may yield fewer and lower quality interactions. Coleman (1988) expresses concern that the job-holding will weaken the "social capital" that depends on the relationships in which children are embedded. These effects may also vary with household characteristics and age of the child. If well-off parents provide higher

[5] Leibowitz (1974a) provides an early example of research highlighting parental time investments in children. There is wide agreement that children benefit from being in households with higher incomes but debate over the nature of the relationship. Duncan & Brooks-Gunn (1997) emphasize the causal effect of poverty, whereas Mayer (1997) suggests weaker income effects, possibly because a greater portion of family income is devoted to children in poor than wealthy families (Lazear & Michael, 1988). All else equal, more time will also be invested in children with few siblings and in two-parent (versus single-parent) households (Haveman & Wolfe, 1995; McLanahan, 1997).

[6] Employed females may "protect" the most productive time by cutting back least on activities that directly engage children. Increased time investments by fathers could offset the reductions of working mothers (in two family households) but little empirical evidence supporting this possibility has yet been provided. The involvement of fathers in child care is rising over time (Bianchi, 2000) but paternal labor supply is unrelated or negatively related to the hours fathers spend with children (Gershuny, 2000; Hofferth, 2001). This last issue receives attention in the analysis of paternal employment below.

quality time, for instance, employment could be more harmful in rich than poor families. However, wealthier families can afford better day care and educated women spend a greater proportion of their nonmarket time in child-related activities (Leibowitz, 1974b), possibly reversing this pattern.[8]

These approaches all stress the beneficial effects of household income and direct parental investments in children, suggesting a potential tradeoff between the two. To illustrate, consider a production function where child outcomes at age t (C_t) depend on status in the previous period (C_{t-1}), the non-market (leisure) time of parents (L), purchased inputs like food or medical care (F), and exogenous determinants or production shocks (V) according to:

(1) $$C_t = C(C_{t-1}, L_t, F_t, V_t).$$[9]

Using subscripts to indicate partial derivatives, C_L and C_F are positive. The production function has several important characteristics. First, parental leisure is good for children. This occurs either through direct time investments or indirectly through reductions in stress, increased energy levels, and so forth. Second, higher incomes raise the ability of parents to purchase productive inputs and influence their time allocation decisions. Third, child outcomes depend partly on prior status and therefore on endowments and the past choices of parents.

Parents have a time constraint

(2) $$L_t + H_t = 1,$$

where H and L are the proportions of total time spent in employment and nonmarket activities. The household also has a budget constraint that limits purchases of child inputs and other consumption to the amount of earned and nonearned income.[10] Solving (2) for H and recursively substituting in for lags of C, equation (1) can be rewritten as:

(1') $$C_t = C(\mathbf{H}, \mathbf{F}, \mathbf{V}),$$

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[7] For instance, working mothers sleep 6 hours (or 10 percent) less per week than their nonemployed counterparts (Bianchi, 2000).

[8] The average quality of child care increases with income, although center-based care for low-income children appears to be of relatively high quality (NICHD Early Child Care Research Network, 1997a). This is probably due to early intervention programs (such as Head Start) available to poor families.

[9] This model follows Becker (1981) in emphasizing the role of non-market time in household production and Grossman (1972) in treating health as an outcome produced by investment activities.

[10] Total time is actually endogenous since it varies with the number of parents and children in the household. The econometric analysis deals with this by including detailed controls for family structure.

where **H**, **F**, and **V** are vectors of current and lagged values (e.g. **H**={H_t, H_{t-1}, … H_{t-n}}, for t-n the first period where parental inputs affect child outcomes). Maximizing C subject to income and time constraints yields the reduced-form demand function:

$$C_t = C(\mathbf{P},\mathbf{V}), \tag{3}$$

where **P** is a vector of current and lagged prices and wages.[11]

Estimates of the child production or reduced-form demand functions specified by equations (1') or (3) provide policy-relevant parameters of structural determinants of child development. However, data restrictions preclude estimation of either model, since information is lacking on the full vector of relevant prices and many important individual-specific production shocks. Instead, the analysis focuses on what Rosenzweig and Schultz (1983) refer to as "hybrid" equations. These take the form:

$$C_t = C(\mathbf{H},\mathbf{X},\varepsilon), \tag{4}$$

where **H** measures work hours, **X** is a vector of individual or family background characteristics, and ε is a disturbance term capturing production shifters or shocks not otherwise controlled for.

A limitation of the hybrid equations is that the coefficient estimates generally embody both the technological properties of the production function and the characteristics of unobserved household preferences or production shifters. For example, the proportion of total income devoted to children could vary with family background characteristics, and parental attitudes about labor supply could be correlated with other investments in children. Child outcomes also generally depend on the *quality* as well as the *quantity* of parental time, and the "technologies" in place when decisions are made. For instance, since child care is one purchased input, the effects of employment partly depend on the difference in the quality of parental and nonparental care, with the result that any negative effects of labor supply are likely to rise as the (price-adjusted) quality of day care falls.[12] A fully specified model would account for the endogeneity between

[11] Formally, parents solve a dynamic programming problem where utility depends on child outcomes, parental consumption and non-market time, and is maximized subject to time and budget constraints. Blau et al. (1996) detail such a model.

[12] This could be modeled by specifying the child production function as $C_t = C(C_{t-1},q_tL_t,F_t,V_t)$, where q is an efficiency parameter reflecting differences in the quality of parental and nonparental care. Parental time could then have a negative effect ($q<0$), if its quality is sufficiently low relative to day care.

these technologies and parental decisions but the regression estimates below do not. Instead, the employment coefficients indicate the "effects" of working given the average differences in other factors (such as child care) that accompany the variation in labor supply. This procedure is reasonable if the time allocation decision precedes and determines the use of these other inputs. It is less appropriate for considering how parental employment might affect child development in a different institutional environment.[13]

These limitations further imply that a causal interpretation can only be applied to the parameter estimates on parental employment if the variables included in **X** capture the effects of all other structural determinants of child outcomes. A particularly important concern is unobserved heterogeneity in family or child characteristics that is correlated with parental job-holding. This receives attention below.

C. Econometric Issues

The model above is operationalized by assuming that the cognitive assessment score for child i at age t (C_{it}) is an additive separable function of parental work hours at child ages t through t-n ($\mathbf{H}_{it}=\{H_{it}, H_{it-1}, \ldots H_{it-n}\}$) and other production shifters (V_{it}), according to:

$$C_{it} = a + \mathbf{H}_{it}\mathbf{b_t} + V_{it} + e_{it}, \quad (5)$$

for e an i.i.d. disturbance.[14] Implicit in (5) is the assumption that parental job-holding prior to t-n or after t has no impact on child outcomes at age t. For this analysis, I assume that t-n is the period immediately following birth.[15]

Several important econometric issues can be clarified using an even simpler (but unrealistic) model where only contemporaneous employment affects child outcomes (i.e. $\mathbf{H}_{it}=\{H_{it}\}$). In the absence of additional controls, this implies the regression equation:

$$C_{it} = \alpha + H_{it}\beta + \varepsilon_{it}, \quad (6)$$

[13] Ermisch and Francesconi (2001) use the alternative approach of estimating a conditional demand function (developed by Pollack, 1969) that includes employment status (to proxy parental time) and other production shifters. Their key assumption is that sibling differences in parental employment during the early years are uncorrelated with child-specific endowments. This assumption is discussed below.

[14] The model could be modified to include lags of the production shifters.

[15] This assumption could be violated if, for example, intertemporal financial transfers cause work before birth to raise consumption during the child's early years.

where $\varepsilon_{it} = V_{it} + e_{it}$. $\hat{\beta}$ is then biased if $cov(H_{it},\varepsilon_{it}) \neq 0$, which occurs if V is correlated with H. For example, there is a spurious positive relationship if employed women have high home productivity or their children have endowments associated with high levels of cognitive skill.

A primary econometric strategy is to use the detailed information in the NLSY to directly account for many potential confounding factors. Thus,

$$C_{it} = \alpha + H_{it}\beta + \mathbf{X}_{it}\gamma + \mu_{it}, \tag{7}$$

consistently estimates the effects of parental employment if the supplementary regressors ($\mathbf{X}$) control for other determinants of child development, such that $cov(\mathbf{X}_{it},\mu_{it}) = cov(H_{it},\mu_{it}) = 0$.

It is useful to decompose the $\mathbf{X}$ into $\mathbf{X}_1$ and $\mathbf{X}_2$, where elements of $\mathbf{X}_1$ *cause* child outcomes while $\mathbf{X}_2$ includes proxies of unobserved determinants that do not themselves have a causal effect. For instance, maternal education would be in $\mathbf{X}_1$ if schooling raises the quality of time spent with children. Conversely, the age at which the mother smoked her first cigarette would fit into $\mathbf{X}_2$ if early tobacco use does not affect child development but instead is correlated with factors (such as maternal discount rates) that are important. This distinction is relevant because the regression coefficients represent parameters of a structural production function if equation (7) is purged of omitted variables bias by including only $\mathbf{X}_1$ variables. However, this analysis incorporates many proxy measures ($\mathbf{X}_2$ variables), so that a causal interpretation can not be applied to the parameter estimates for these regressors. For instance, early smoking may be negatively correlated with child outcomes without having a causal effect.

The goal of this analysis is to estimate the impact of parental employment by including a sufficiently rich set of covariates that the error term is orthogonal to $\mathbf{H}_{it}$. However, in doing so, it is important to exclude variables that *result* from parental job-holding (such as the home environment during the child's early years), since these capture a portion of the labor supply effect. One exception is that some models control for family income or the use of nonparental child care, both of which are likely to be affected by work hours, in order to examine the extent to which changes in these factors explain a portion of the effect of parental employment.

Most econometric models also hold constant maternal employment after the child's third year. Such controls have rarely been included in previous research. One result is that the

estimated effects are likely to have combined the impact of working during the first three years with those of labor supply in later periods. To illustrate, consider the case where $\mathbf{H}_{it}=\{H_{it}, H_{it-1}\}$, for H_{it} corresponding to employment at the assessment period and H_{it-1} to labor supply in the early childhood period of interest. Assume that regressions of:

$$C_{it} = \alpha + H_{it-1}\beta_{t-1} + H_{it}\beta_t + \mathbf{X}_{it}\gamma + \mu_{it} \tag{8}$$

would yield unbiased estimates but that the model estimated is

$$C_{it} = \alpha + H_{it-1}\beta_{t-1} + \mathbf{X}_{it}\gamma + \lambda_{it}, \tag{8'}$$

where $\lambda_{it} = H_{it}\beta_t + \mu_{it}$. It is straight-forward to show that $\text{cov}(H_{it-1},\lambda_{it}) = \beta_t \text{ cov}(H_{it-1},H_{it})$, implying that $\hat{\beta}_{t-1}$ is likely to be biased in the direction of β_t since labor supply is positively correlated over time.[16]

Even an extensive set of explanatory variables may not fully account for all important sources of heterogeneity. However, under reasonable assumptions, employment *prior* to birth can be added to the regression equation to test for bias due to unobserved maternal factors.[17] To see this, decompose the error term in (7) as $\mu_{it} = M_i + K_i + \varphi_{it}$, where M are time-invariant maternal characteristics, K are child-specific endowments (defined such that $\text{cov}(M_i,K_i) = 0$), φ is an i.i.d. error, and allow H to be related to M and K according to:

$$H_{it} = c_tM_i + d_tK_i + \pi_{it}. \tag{9}$$

If π is an i.i.d. disturbance term,

$$\text{cov}(C_{it},H_{it}) = \beta \text{ var}(H_{it}) + c_t \text{ var}(M_i) + d_t \text{ var}(K_i) \tag{10}$$

and the last two terms in (10) indicate the biases resulting from the failure to control for omitted maternal and child effects.

If work hours in the year prior to birth (H_{it-p}) have no effect on child outcomes,

$$\text{cov}(C_{it},H_{it-p}|H_{it}) = c_{t-p}\text{var}(M_i), \tag{11}$$

implying that in the augmented regression model

[16] The correlation between average hours in years 2 and 3 and those in year 4 is .754 in the representative NLSY subsample. I control for maternal employment in year 4 but not in years 5 and 6, when analyzing the reading and math scores of 5 and 6 year olds. One implication is that the coefficient on year 4 employment partially captures the effects of work performed after this date.

[17] This approach draws heavily on Gottschalk's (1996) research examining the intergenerational correlation in welfare participation.

(12) $$C_{it} = \alpha + H_{it}\beta + \mathbf{X}_{it}\gamma + H_{it-p}\delta + \mu_{it},$$

$\hat{\delta}$ provides an estimate of the heterogeneity bias due to the unobserved mother-specific effects. In particular, if $c_t = c_{t-p}$, $\hat{\delta}=0$ implies that this source heterogeneity has been completely accounted for.

The procedure just discussed will generally *not* test for bias due to unobserved child-specific factors. This is even more of an issue in recent research that includes maternal fixed-effects as alternative method of controlling for heterogeneity in family backgrounds (e.g. James-Burdumy, 1999; Neidell, 2000; Ermisch and Francesconi, 2001; Waldfogel et al., forthcoming). The fixed-effect (FE) models correspond to:

(13) $$C_{iqt} = \alpha + H_{iqt}\beta + X_{iqt}\gamma + M_q\delta + \phi_{iqt},$$

where C_{iqt} is the assessment score at age t of child i with mother q, M is a vector of mother-specific dummy variables, and $\phi_{iqt} = K_i + \varphi_{iqt}$. Such specifications examine how sibling test scores differ as a function of maternal employment at given child ages. This can be seen by representing sibling differences in cognitive scores, maternal employment, and personal characteristics using ΔC, ΔH, and $\Delta\mathbf{X}$ and rewriting (13) as:

(14) $$\Delta C_{it} = \alpha + \Delta H_{it}\beta + \Delta\mathbf{X}_{it}\gamma + \Delta\phi_{it}.$$

Notice that the maternal fixed-effect has been differenced away, eliminating unobserved heterogeneity from this source. However, child-specific endowments have not been controlled for and there is no reason to believe that the FE models will provide consistent estimates or even reduce the size of the bias. Specifically, since the error term in (14) is $\Delta\phi_{it} = \Delta K_i + \Delta\varphi_{it}$, the coefficients are biased if $\text{cov}(\Delta H_i, \Delta K_i) \neq 0$ and the bias may be *larger* than in corresponding OLS models if unobserved child-specific factors are a key determinant of sibling differences in maternal labor supply.[18] This is important because there is a good deal of evidence that mothers work less when children have health or developmental problems.[19] In this case, $\text{cov}(\Delta H_{it}, \Delta K_i) > 0$ and the FE models will understate (overestimate) the costs (benefits) of maternal employment.

[18] The sample analyzed in fixed-effect models is also less representative since it is restricted to siblings in the age groups where complete data are available.

[19] Behrman, Pollack, and Taubman (1982) provide early research showing compensatory investments by parents in siblings with lesser genetic endowments. More recently, Neidell (2000) finds that low birthweight or slow developmental abilities delay the return of mothers to work; Ermisch and Francesconi

D. Data and Descriptive Results

Data are from the National Longitudinal Survey of Youth, a sample of U.S. residents born between January 1, 1957 and December 31, 1964 and who have been surveyed since 1979.[20] In 1982 the NLSY began including questions on pregnancy, pre-natal, and post-natal care. Children born to and living with female NLSY respondents have been interviewed at two year intervals beginning in 1986, with information used below through 1996. The combined data set provides a unique source of longitudinal information on a large sample of children and their parents, including great detail on maternal, child, and household characteristics.

The NLSY (through 1996) supplies data on children whose mothers were between the ages of 29 and 38 at the end of 1995. This covers approximately 80 percent of childbearing for this cohort but is not fully representative of all fertility, since the offspring of some older women (who tend to have high incomes and education levels) are excluded. Although this selection should be kept in mind when interpreting the results, the sample is *more* representative than in most of the previous studies summarized in Table 1.

Cognitive development is proxied by scores on the Peabody Picture Vocabulary Test (PPVT) and the Peabody Individual Achievement Test Reading Recognition (PIAT-R) and Mathematics (PIAT-M) subtests.[21] These are among the most widely used assessments of preschool and early school-aged children and are known to have high test-retest reliability and concurrent validity (Baker et al., 1993). The PPVT measures receptive vocabulary for Standard American English and provides a quick estimate of verbal ability and scholastic aptitude. The PIAT-R indicates word recognition and pronunciation ability by examining skills such as matching letters, naming names, and reading single words aloud. The PIAT-M assesses

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cite evidence of reduced parental employment when children have limiting health conditions or disabilities; and Norberg (1998) shows that mothers of "high risk" children return to jobs relatively slowly, as do those whose children have low development scores or high levels of "fearfulness".

[20] The NLSY originally included a representative sample of 6,111 youths, an oversample of 5,295 blacks, Hispanics and economically disadvantaged whites, and a supplemental sample of 1,280 persons in the military in September 1978. Interviews with the military subsample were suspended after 1984 and for economically disadvantaged non-Hispanic whites after 1990. This data set is now sometimes referred to as the NLSY79, to distinguish it from the new NLSY97 survey which covers a later cohort.

[21] The NLSY uses a revised version of the original PPVT test frequently referred to as the PPVT-R.

attainment in mathematics beginning with early skills, such as recognizing numerals, and progressing to advanced concepts in geometry and trigonometry.

The PPVT was administered to children aged 3 and over in 1986, with additional assessments for previously untested age-eligible children in 1988, 1990, 1992, 1994, and 1996.[22] This investigation examines results for 3 and 4 year olds. The PIAT-M and PIAT-R were given to children 5 and over in each survey year, with data used below on 5 and 6 year olds.[23] The analysis focuses on the "standard" assessment scores. These represent transformations, on an age-specific basis, of the raw scores that were originally (during the 1970s) designed to have a normal distribution with a mean of 100 and a standard deviation of 15. Standard scores have been commonly used by previous researchers (e.g. Baydar & Brooks-Gunn, 1991; Blau & Grossberg, 1992, Parcel & Menaghan, 1994). Models are also estimated using raw or percentile scores to insure that the results are not sensitive this choice.[24] For ease of interpretation, the dependent variables have been normalized to have a standard deviation of one. Therefore, the regression coefficients show the standard deviation change in the test score predicted by a one unit change in the explanatory variable.

Maternal employment is measured in the year prior to birth (denoted as year 0) and in the first through fourth years of the child's life (years 1 through 4). Most regressions control for average weekly work hours divided by 20.[25] A one unit change thus corresponds to switching from no employment to half time work or from 20 to 40 hours per week on the job. To check the robustness of these results, some models instead control for: the proportion of weeks worked

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[22] All age-eligible children were assessed in 1992, even if they had previously been tested.

[23] Reading comprehension is also assessed for children receiving sufficiently high PIAT-R scores. However, this test rarely provides useful information for children under 7 and it is not utilized here.

[24] The percentile scores are derived from the standard scores and represent a uniformly-distributed ranking of individuals in the original age-normed distribution. The raw scores are closer to a log-normal distribution, particularly for the PIAT-R and PIAT-M assessments. One implication is that the percentile scores will tend to place more weight on observations in the middle of the distribution and raw scores on those at its upper-tail than will the standard scores.

[25] Work hours cover all jobs held by the respondent. In the few cases where information on secondary jobs is missing, hours are calculated using data for the main job only. NLSY mothers with a job but on maternity leave are generally categorized as employed, implying that employment rates immediately after birth are overstated and the effects of maternal work during this period may be underestimated. However, few women are on maternity leave for more than a few weeks so the resulting bias is likely to be small.

(multiplied by two); whether the mother worked at all during the specified period; or the number of weeks after birth until she resumes employment.[26]

The NLSY contains limited information on the employment of fathers residing with interviewed mothers. Specifically, data for the *calendar year* preceding the survey date are used to control for average weekly employment hours or weeks worked during the year.[27] The analysis of paternal employment is restricted to children with fathers living in the household at the interview date of the fourth calendar year after birth, since these men are likely to have been with the child throughout the first three years. For children born before (on or after) July 1, year 1 refers to the calendar year of birth (the next calendar year). Thus, year 1 includes the majority of the child's first year of life. Years 0, 2, 3, and 4 are defined in reference to year 1.[28]

This analysis exploits the extensive child, maternal, and household information available in the NLSY. A vector of "basic" background variables, so labeled because they have frequently been controlled for in previous research, contains continuous measures of birth order, mother's age at child birth (in years), her highest grade completed, and a quadratic for child age (in months). Also included are dummy variables for race/ethnicity (2 variables), residence in an SMSA or central city (2 variables), and sex of the child. Unless noted, all regressors are measured at the child assessment date. Table A.1 provides detailed descriptions of the variables used in this analysis.

Most models also control for a set of "supplemental" characteristics such as: the mother's Armed Forces Qualifications Test (AFQT) score, her marital status at the survey date before pregnancy and the next 3 surveys (8 variables), and whether family income was below the poverty line in the second calendar year before the child's birth.[29] Some form of these variables have occasionally been used by prior researchers. The vector also contains maternal, family, and

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For instance, Klerman & Leibowitz (1994) find that just 2.2 percent of mothers were on paid leave and 6.0 percent on unpaid leave three months after giving birth during the late 1980s.

[26] The proportion of weeks is multiplied by two so that a one unit change corresponds to switching from no work to employment in half the weeks, analogous to a one unit change in the hours variable.

[27] The more detailed employment history included for mothers is not available for spouses.

[28] Father's weeks worked are not reliably reported prior to 1981 (there are virtually no observations with zero weeks.) Therefore, data on paternal employment are restricted to the period after 1980.

[29] Poverty in the second year before birth is used to avoid confounding caused by any changes in employment during pregnancy or after birth.

child factors that may directly influence or indirectly proxy determinants of child development but have not previously been controlled for. These include: the mother's number of siblings, her location at age 14 (3 variables), whether magazines, newspapers, or library cards were in her home at 14 (3 variables), the place of birth and education of her parents (4 variables), her family structure at age 14, whether she attended a public or private secondary school, if the child's father lived in the home at the assessment date, and relative ages of the child's siblings (4 variables).

The supplemental covariates also account for maternal attitudes and experiences that may be related to child outcomes (either directly or as proxies for unobserved causal factors) such as: the mother's religious affiliation and church attendance in 1979 (6 variables), age at which she smoked her first cigarette (4 variables), drug use prior to age 21 (4 variables), and her cumulative score on a family roles attitude scale.[30] Early child health or developmental problems and pregnancy characteristics are also incorporated through indicators of: low weight and premature births (7 variables), long hospital stay at birth, hospitalization during the first year of life, physician visits related to illness in the first 12 months, excessive or deficient weight gain and smoking or alcohol use by the mother during pregnancy (5 variables), and her height-adjusted weight prior to pregnancy (4 variables).[31]

A third set of regressors, labeled "maternal employment characteristics", control for maternal work hours in year 4, as well as the wages and occupation of the mother immediately prior to pregnancy (12 variables). These covariates supply information on the opportunity costs of not working and may be correlated with unobserved parental influences on child development. Finally, some models hold constant family incomes, the use or type of day care, and the age and highest grade completed by the father in the calendar year of birth.

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[30] The scale was constructed by summing over the eight questions, with each assigned a value of 0 or 1. Higher scores indicate greater work orientation and less traditional attitudes. For instance, one was added to the scale for women who *disagreed* or *strongly disagreed* with the statement "a wife who carries out her full family responsibilities doesn't have time for outside employment", as well as for those who *agreed* or *strongly agreed* that "men should share the work around the house with women, such as doing dishes, cleaning, and so fourth".

[31] Height-adjusted weight is categorized by body mass index (BMI), defined as weight in kilograms divided by height in meters squared. Height is measured in 1981 and weight immediately before pregnancy. Maternal employment could influence the probability of the child being hospitalized during the first year, requiring caution when using this regressor. However, virtually identical results are obtained when this variable is excluded.

Data on one or more background characteristics are missing for some respondents. To avoid excluding these persons, the relevant regressors are sometimes set to zero and dummy variables created denoting the presence of missing values. For example, mothers not reporting an AFQT score are given a value of zero and the "missing AFQT" variable is set to one.[32] Alternatively, some dummy variables are valued at one when the specified condition is met and zero when *either* it is not *or* the relevant data are missing.[33]

D.1 Patterns of Maternal Employment

Figures 1 through 5 provide descriptive information on maternal employment during pregnancy (the three quarters prior to birth) and the first three years of the child's life, using data from the nationally representative subsample of the NLSY.[34] Figure 1 displays histograms for weeks worked and shows that there is always substantial bunching at the extremes of employment in all or no weeks – accounting for 55 percent of pregnant women, 54 percent of mothers during year 1, and 63 percent in years 2 and 3. While there is somewhat more dispersion in work hours, Figure 2 demonstrates that most women either do not hold jobs or work quite intensively. For instance, 66 percent are not employed or average more than 30 hours per week during pregnancy, as do 64, 66, and 66 percent in the first, second, and third years of the child's life.[35]

The concentration of employment is more sharply highlighted in Figure 3, which calculates average hours in weeks of work. Conditional on some employment, between 26 and 35 percent of women work exactly 40 hours and 67 to 74 percent average 30 or more hours. By contrast, fewer than 1 in 4 (1 in 6) are employed less than 25 (20) hours per week. Thus, most mothers with young children either do not hold jobs or work close to full-time, suggesting that the econometric analysis is likely to obtain similar results whether controlling for average work hours or the proportion of weeks worked.

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[32] This was done for number of siblings, marital status, age of smoking initiation, location and language spoken in the home at 14, presence of the father in the household, poverty status before birth, education of the mother's parents, birth weight, and gestational age.

[33] For example, this strategy was used for hospitalizations and doctor visits in the first year, pregnancy behaviors, and residence in an SMSA at the survey date.

[34] Similar results are obtained using weighted data for the full sample.

[35] Work hours are top-coded at 50 in the histograms for expositional convenience.

The "unconditional" probabilities in Figure 4 indicate that 72 percent of mothers work during pregnancy but that less than one-fourth (24 percent) do so until giving birth. The "conditional" estimates show employment probabilities for those with some pregnancy work experience and demonstrate that one-third of such women remain employed until delivery. The figure also illustrates that weekly hazard rates out of employment (displayed in percent terms and multiplied by 5) are less than 2 percent per week during the first two-trimesters, increase to around 3 percent in the seventh month, 4 to 5 percent in month 8, and then rise to almost 20 percent in the week before delivery.

Figure 5 summarizes reemployment rates following birth. One-fourth of mothers are absent from jobs for less than one week after delivery, although some "employed" women may initially be on maternity leave. Weekly reemployment hazard rates average 4 to 8 percent for the next 10 weeks, decline rapidly to 1 to 2 percent in the fourth through sixth months of the child's life, and are less than 1 percent for the remainder of the first year. The hazard rates typically range between 0.2 and 0.6 percent during the second year and 0.1 to 0.2 percent in year 3. Sixty-five, 77, and 84 percent of women return to jobs prior to their child's first, second, and third birthday, as do 53, 70, and 79 percent of those taking some time off work after birth.

D.2 Paternal Employment

Most mothers remain home with infants for a substantial period of time. The same is not true for fathers. As detailed in the top half of Figure 6, 64 percent of men work in every week of year 1, 84 percent are employed at least four-fifths of weeks, only 3 percent do not hold a job at any point, and just 8 percent work in less than 26 weeks.[36] Similarly, 41 percent of fathers are employed every week of the first three years, 84 percent in 80 percent of weeks, less than 1 percent never hold a job, and fewer than 4 percent work in less than half the weeks.

There are at least four reasons to believe that most weeks away from jobs do *not* occur because fathers are choosing to spend time with young children. First, nonemployment is evenly distributed across the first three years, rather than being concentrated in the year 1, as expected if

[36] Information in this section refers to fathers living with mothers in the nationally representative subsample of the NLSY at the survey date of the fourth year after birth.

the absences are motivated by a desire to be with infants.[37] Second, over half (53 percent) of nonemployment is spent on temporary layoff or looking for work, suggesting that most work absences are involuntary.[38] Third, fathers with substantial joblessness possess characteristics associated with employment instability, so that their nonemployment may be due to these factors.[39] Fourth, time-diary evidence presented by Hofferth (2001) indicates that nonworking fathers spend *less* time with their children than employed men.

The bottom half of Figure 6 shows the distribution of paternal hours in weeks of employment.[40] Once again, there is substantial bunching around 40 hours per week – over half of men work exactly this amount in year 1. However, in sharp contrast to mothers, a large proportion of fathers are on jobs for considerably longer hours: 38, 29, and 12 percent average at least 45, 50, and 60 hours of work per week in year 1; 41, 23, and 8 percent do so throughout the first three years. Conversely, only 10 percent of mothers average 45 or more hours in weeks of employment during the child's first year.

D.3 Descriptive Relationships

Early maternal employment is associated with relatively *high* levels of child cognitive achievement. The top panel of Table 2 shows that children whose mothers worked 30 or more hours per week during the first three years scored a statistically significant .2 to .3 standard deviations higher on the PPVT, PIAT-R, and PIAT-M tests than those with nonworking mothers.

The remainder of the table demonstrates that these differences are probably *not* primarily due to a causal effect of the labor supply. Instead, children with nonemployed mothers come from disadvantaged families and possess characteristics associated with delayed development. For instance, women not holding jobs in the first year are less educated and have lower AFQT scores than those working 30 or more hours per week. They are also more likely to be poor

[37] Among fathers jobless for less than six months during the three years (84 percent of the sample), 34, 35, and 31 percent of the weeks take place during years 1, 2, and 3. By contrast, 63 percent of the weeks corresponding mothers are off work occur during the first year and just 22 percent in year 3.
[38] A larger proportion of nonemployment is devoted to job search or spent on layoff in year 1 than in years 2 and 3, which would not be expected if fathers were choosing to invest time in infants.
[39] For instance, 28 percent of fathers jobless at least six months during the three years had less than 12 years of education, compared to 11 percent of those out of work fewer than 26 weeks; 28 percent of the former group were under the age of 25 versus 19 percent of the latter.
[40] Hours are top-coded at 70 in the histograms for expositional convenience.

before pregnancy (30 vs. 6 percent), unmarried at childbirth (58 vs. 70 percent), and to have the father absent from the household at the survey date of the fourth calendar year after birth (29 vs. 20 percent). Similarly, the children of these mothers more often have low birthweight and require hospitalization before their first birthday, both of which proxy poor health.

These findings indicate the need for a careful multivariate investigation. If disadvantaged family backgrounds or child health problems retard cognitive development and are associated with reduced labor supply, failing to correct for this heterogeneity will lead to upwards biased estimates of the return to parental employment. The conclusions of many previous analyses, that include only rudimentary controls, are therefore likely to be overly optimistic.

E. Regression Estimates

One empirical strategy is to examine whether the addition of more complete controls for confounding factors lowers the parameter estimates on maternal employment. A second is to pay special attention to the coefficient on work hours in the year prior to birth. As discussed, a small and statistically insignificant parameter estimate for this variable indicates that the other covariates may adequately account for nonrandom selection into maternal employment, whereas a large or statistically significant coefficient suggests remaining omitted variables bias.

Table 3 displays the findings of nine specifications for each of the three cognitive assessments. The dependent variables in models (a) through (f) and (i) are standard test scores, normalized to have a standard deviation of one. Columns (g) and (h) show results for (similarly normalized) percentile and raw scores. Maternal employment refers to average weekly work hours (divided by 20) in the specified period, except for model (h) which examines the fraction of weeks worked (multiplied by 2). Year 0 includes the four quarters prior to birth, year 1, 2, and 3 to the 1st through 4th, 5th through 8th, and 9th through 12th quarters after it. All models include assessment year dummy variables. Additional regressors are detailed at the bottom of the table: B, S, and E refer to the vectors of basic, supplemental, and maternal employment characteristics described previously and detailed in Table A.1; I indicates controls for total family income ($1996) during the calendar year before birth and the next 4 years.[41]

[41] This includes income from jobs, business/farm activities, government transfers (AFDC, Food Stamps, SSI/public assistance, unemployment insurance, veteran's/disability benefits), alimony, and child support.

E.1 Verbal Skills of 3 and 4 Year Olds

The top panel of Table 3 summarizes the results for PPVT scores. As in the descriptive analysis, 3 and 4 year olds with employed mothers have relatively high verbal ability. In column (a), which controls only for work hours and the assessment year, 20 hours per week of labor supply during the child's first (second and third) year is associated with a .05 (.10) standard deviation rise in verbal performance.

This positive relationship, largely results from omitted variables bias, rather than any causal effect of maternal job-holding. Thus, the inclusion of the basic set of covariates cuts the parameter estimate for years 2 and 3 by more than 40 percent and switches the year 1 coefficient from positive to negative (see model b). These results closely resemble those of many previous researchers using similar models in suggesting harmful "effects" of maternal employment during the first year but with roughly offsetting benefits for working during the next two years. However, the inclusion of the supplemental and maternal employment characteristics further reduces the predicted benefit of working in years 2 and 3 (columns c and d), and holding constant work hours in the year prior to birth substantially strengthens the negative effect found for labor supply in year 1. In specification (e), a 20 hour per week increase in employment throughout the first three years is correlated with a .04 standard deviation decline in PPVT scores, suggesting considerably more deleterious effects then are obtained when including less comprehensive regression controls. The year 0 coefficient is substantial (.050) and estimated fairly precisely, suggesting that even the extensive set of explanatory variables may not adequately account for heterogeneity. I discuss this in greater detail below.

Maternal employment could benefit children by raising earnings. However, as shown in specification (f), holding family income constant has little effect on the hours coefficients. One explanation is that work may be associated with decreases in other sources of financial support (e.g. transfer payments or spousal earnings) so that incomes actually do not rise much. A second is that the direct income effect, estimated from the regressions, is extremely small.[42]

[42] A $10,000 increase in annual income beginning in the year before birth and sustained for the next four years raises predicted PPVT scores by just .003 standard deviations.

The negative impact predicted of job-holding in year 1 is somewhat greater when using percentile rather than standard scores (specification g) and that for employment in years 2 and 3 is slightly smaller for the raw scores (column h), but neither difference is large or significant and the coefficient on pre-birth work hours suggests that heterogeneity may be less well controlled for in these models. Holding weeks rather than hours of work constant (model i) attenuates the estimated effects (as expected since some hours variation occurs during weeks of employment) but only slightly. Thus, the predicted impact of maternal employment is relatively robust to these changes in specification.

E.2 Are the Results Consistent Across Alternative Cognitive Assessments?

The findings for PPVT scores suggest that previous research may present an overly optimistic evaluation of the effects of early maternal employment to the extent that it fails to adequately account for confounding factors. This section demonstrates that the relationship between job-holding and child development is also sensitive to the age of testing or types of skills assessed, with more negative results obtained for the reading and mathematics achievement of 5 and 6 year olds.

PIAT-R and PIAT-M scores are the dependent variables in the middle and bottom panels of Table 3. The inclusion of more complete controls once again yields less favorable predicted impacts of maternal employment. Absent covariates other than the survey year, there is a positive association between work and child achievement (column a). This correlation shrinks and loses statistical significance or becomes negative when the "basic" regressors in (column b) are included, declines further with the addition of the supplemental characteristics (specification c), and decreases still more when maternal employment characteristics and pre-birth work hours are held constant (columns d and e). The inclusion of family incomes does not materially affect the results (specification f), and broadly similar parameter estimates are obtained when percentile or raw scores are the dependent variables (models g through h), or when weeks rather than hours worked are controlled for (column i).

These similarities notwithstanding, the expected impact of early employment is markedly more negative for PIAT-R and PIAT-M achievement than for PPVT scores. Most strikingly,

working during years 2 and 3 is predicted to have strong detrimental impacts, in contrast with the small positive correlation for the PPVT test. In specification (e), 20 hours of additional work per week during the second and third years is associated with statistically significant .08 and .06 standard deviation reductions in reading and mathematics scores. An extra 20 hours per week of labor supply throughout the first three years lowers expected PIAT-R and PIAT-M scores by .11 and .09 standard deviations, versus a smaller .04 standard deviation predicted decrease in PPVT performance. These correspond to drops in test scores from the median to the 46^{th}, 47^{th}, and 49^{th} percentiles respectively.

E.3 Remaining Heterogeneity

The usefulness of the pre-birth employment coefficient as a test for unobserved heterogeneity was confirmed by estimating models equivalent to columns (a) through (d) in Table 3 but with the inclusion of year 0 work hours. As anticipated, the magnitude and statistical significance of the test coefficient decline with the inclusion of additional controls.[43] However, these additional estimates did not reveal any consistent method of purging the year 1 through 3 coefficient of remaining heterogeneity.[44] Therefore, I adopt the limited strategy of using the pre-birth parameter estimate to test for omitted variables bias without attempting to further correct for it. As mentioned, the test coefficient in the preferred model (specification e) is relatively large and close to significance for PPVT scores. By contrast, the smaller magnitude suggests that uncontrolled heterogeneity is less of an issue for the PIAT-R or PIAT-M tests.

[43] The coefficients (standard errors) are .301, .074, .053, and .050 (.026, .023, .028, and .028) for PPVT scores in models corresponding to specifications (a), (b), (c) and (d). For PIAT-R performance they are .234, .081, .059, and .030 (.024, .024, .023 and .029) and for the PIAT-M test .233, .079, .062, and .037 (.024, .024, .024 and .030). Large positive parameter estimates in models with limited covariates provide further evidence that children with working mothers come from advantaged family backgrounds. A potential issue is that work during pregnancy might affect cognitive development. As an alternative specification test, I therefore estimated models with separate controls for employment in each of the four quarters prior to birth. Notably, the coefficient on work hours in the fourth quarter before birth (i.e. prior to pregnancy) was generally positive and large in specifications with limited controls but declined in size as heterogeneity was more completely accounted for.

[44] In the notation of equation 12), the additional covariates sometimes *increase* $\hat{\beta}$ by approximately the same amount $\hat{\delta}$ declines. In other cases $\hat{\beta}$ changes little or decreases by roughly the same magnitude. These three scenarios alternatively suggest using $\hat{\beta}-\hat{\delta}$, $\hat{\beta}$, or $\hat{\beta}+\hat{\delta}$ to provide heterogeneity-corrected estimates of the effect of maternal employment.

E.4 Intensity, Timing and Interaction Effects of Early Maternal Employment

Some previous researchers (e.g. Waldfogel et al., forthcoming) focus on the distinction between any versus no work during the child's early years. For comparison, the top panel of Table 4 provides results where average work hours are replaced by dummy variables indicating any (versus no) maternal employment during the specified period.[45] These models distinguish working from not working but do not examine differences in the intensity among mothers holding jobs. The results suggest benefits of specializing in home production during the child's first year and of engaging in at least some employment during the next two.

The second panel allows for nonlinearities by separating part-time and full-time jobs, using 30 hours per week as the threshold between the two.[46] The estimates suggest that part-time employment yields larger benefits or lower costs than full-time work. Part-time jobs during the first year are associated with .04, .06, and .06 standard deviation reductions in PPVT, PIAT-R, and PIAT-M scores. The PPVT and PIAT-R assessments are predicted to be an *additional* .04 standard deviations lower for women with full-time jobs (with no difference for the PIAT-M test). The disparities are even sharper in years 2 and 3. Part-time work is correlated with substantial increases in all three cognitive measures – by .08, .04, and .07 standard deviations – whereas moving from part-time to full-time employment reduces the expected scores by .04,.09, and .11 standard deviations.

An important decision for mothers is how soon to resume employment after giving birth. This is addressed in the third panel of Table 4, by replacing work hours with dummy variables indicating the timing of the return to work.[47] The estimates suggest that cognitive development is enhanced by nonmarket activities during the first year of life and that investments extending into the second and third year may also improve test performance. Compared to resuming work within six months, staying at home for through at least a portion of the second year is correlated with .12, .13, and .09 standard deviation increases in PPVT, PIAT-R and PIAT-M scores.

[45] Except where noted, the remaining regression specifications correspond to column (e) of Table 3.
[46] Similar results are obtained using 25 or 35 hours per week as the cutoff.
[47] Women returning to jobs within six months are the reference group. Work hours during year 4 are deleted from these models, given the inclusion of a dummy variable indicating return to work in the fourth year or later.

Equally large gains in math and reading achievement are anticipated for mothers out of market work for two to three years, although still longer absences yield smaller expected benefits.[48]

The last panel allows the effects of employment during years 2 and 3 to differ depending on whether or not the mother worked in the first year. The results imply that labor supply in these later years may have more favorable effects if no job was held during child's infancy. For example, working 20 hours per week in years 2 and 3 raises predicted PPVT scores by .04 standard deviations if the mother did not hold a job in the first year but by less than two-thirds as much if she did. The PIAT-R assessments of 5 and 6 year olds similarly are expected to fall by .06 standard deviations conditional on no job in year 1, with no effect predicted for PIAT-M performance, whereas these test scores are anticipated to decline by .09 and .08 standard deviations if the mother did work in the first year.

These results modify our understanding of the effects of maternal employment in two ways. First, the uniformly negative predicted effects of labor supply during year 1 emphasize the special importance of investments during the child's first year. The findings also cast doubt on the possibility that infants suffer losses when their mothers work but "catch-up" if the employment continues for the next two years. Second, the evidence suggests that limited employment during the child's second and third is associated with higher test scores but that the benefits decline and become negative as work hours increase. A reasonable overall interpretation is that child cognitive development is maximized when the mother stays home during the entire first year and works part-time during some or all of the next two.

E.5 Subsamples

Table 5 provides results for population subgroups. The specifications are identical to model (e) of Table 3, except that the first column for each outcome shows the parameter estimate for work hours averaged over the entire first three years. The second column displays the coefficient on pre-birth employment, an indication of potential remaining omitted variables bias.

[48] Similar models that control for quadratics in weeks after birth until the mother returns to a job indicate that PPVT, PIAT-R and PIAT-R scores reach a maximum when employment resumes after 95, 105, and 120 weeks. The corresponding increase in predicted test performance is.08, .11, and .06 standard deviations.

Findings for the full sample, displayed in the top panel, confirm that maternal job-holding has a small and statistically insignificant negative expected effect on PPVT scores but larger deleterious predicted impacts on PIAT-R and PIAT-M performance: 20 extra hours per week of work throughout the three years is associated with .03, .11, and .09 standard deviation drops in the three cognitive measures, corresponding to declines from the median to the 49th, 46th, and 47th percentiles.

Some previous researchers have found stronger negative effects for boys than girls (e.g. Desai, et al., 1989; Brooks-Gunn, et al., forthcoming). The second panel tests for, and largely rejects, this possibility. Maternal employment during the first three years is predicted to lower the PPVT and PIAT-M scores of boys by more than girls but the reverse is true for the PIAT-R test and none of the gender differences approach statistical significance.[49]

The third panel provides separate results for (non-Hispanic) whites and blacks. Some recent studies (e.g. Han, et al., 2000; Waldfogel, et al., forthcoming) suggest that the deleterious impacts of early job-holding are largely restricted to white children. However, no systematic differences are observed in this analysis. More negative effects on PPVT scores are predicted for whites but the pattern is reversed for the PIAT-R and PIAT-M assessments and none of the race differences are statistically significant.[50]

The fourth panel divides the sample by whether or not a spouse (not necessarily the biological father) is in the household during *all* or *none* of the survey dates in the three calendar years following the child's birth.[51] Given the relatively small sample sizes, the employment coefficients are estimated imprecisely. Nevertheless, the results hint that maternal job-holding may be more harmful for two-parent than female-headed households. Hill et al. (2001) and Brooks-Gunn et al. (forthcoming) similarly find more negative employment effects for children

[49] Han et al. (2001) and Hill et al. (2001) similarly fail to uncover any gender differences in the impact of employment in the first year; Waldfogel et al. (forthcoming) find larger negative effects of full-time work for girls than boys.

[50] Hill et al. (2001) similarly fail to find any race differences once employment selection is carefully controlled for. The maternal employment coefficients obtained for Hispanic children are: -.035, -.180, and -.118 for PPVT, PIAT-R, and PIAT-M scores. However, the standard errors are also large (.082, .069, and .070), and the sizable year 0 coefficients (ranging from .090 to.145) suggest considerable potential for heterogeneity bias.

[51] Children with a male adult present in some but not all years are excluded from this analysis.

of married than unmarried women. Possible explanations are that the home environment or parental time investments are of relatively high quality in two-adult families or that the earnings of mothers are especially beneficial in female-headed households. However, the coefficients on pre-birth work hours are also bigger for the nontraditional families, suggesting greater potential for omitted variables bias.[52]

The last panel divides the sample into employed women earning more or less than $10 per hour ($1996) in the fourth quarter before birth.[53] Maternal labor supply is negatively related to the PPVT scores of high but not low earners, as expected if the mother's earnings yield relatively large benefits when wages are low. However, a similar result is not obtained for the PIAT-R or PIAT-M assessments.

E.6 Child Care

Nonparental child care has been ignored until now out of concern that its use is influenced by work decisions, so that controlling for it would inappropriately attenuate the effects of maternal employment.[54] However, the impact of job-holding could also vary with day care arrangements. For instance, the deleterious consequences of working during the child's first year might reflect the low average quality of day care in the United States.[55] Some information on these issues was obtained by estimating models with interactions between work hours and the use or mode of nonparental care.[56] The restricted nature of this analysis is dictated by the limited

[52] I also estimated models for subsamples divided by whether or not the biological father was in the household in the fourth calendar year after birth. These specifications indicated larger negative effects of maternal employment on PPVT and PIAT-M scores in families where the father was present, with no differences in PIAT-R performance.

[53] Women not working at this time are excluded. Other wage thresholds were also considered.

[54] Over 85 percent of NLSY mothers with children under the age of three and working at least 30 hours per week use nonparental care, compared to less than one-fifth of corresponding nonemployed women. Day care rises with child age, from 44 percent in year 1 to 54 percent in year 3, because nonworking mothers more often place toddlers (than infants) in care and women with older children are more likely to work. The increase between the first and third year is entirely accounted for by growth in center-based care.

[55] Helburn & Howes (1996) indicate that 86 percent of day care centers provide "mediocre or poor" services and that only 9 percent of family child care homes supply "good" quality care.

[56] The modes include relative, center-based, and other types of nonparental care. The regression model is: $C_{it} = \alpha + X_{it}\beta + H_{it-j}\gamma + H_{it-j}D_{it-j}\delta + \varepsilon_{it}$, where D_{it-j} indicates whether nonparental care is used or the type of the first arrangement.

child care data in the NLSY – there is information on the number and type of arrangements but not (after 1989) on the intensity, cost, or quality of care.

The regression estimates (not shown) fail to uncover any consistent effect of child care. Children placed in care during the first year have slightly lower predicted verbal ability at ages 3 and 4 but marginally higher reading or mathematics achievement two years later, while nonparental care in years 2 and 3 is associated with higher (lower) PPVT and PIAT-M (PIAT-R) scores. None of these "effects" are close to being statistically significant. The data also hint that center-based care during years 2 and 3 may yield some benefits but again the estimates are imprecise. Generally, these results conform to prior research findings of small and inconsistent effects of nonparental care. Most importantly for this analysis, the coefficients on maternal employment are insensitive to the inclusion of controls for day care.

E.7 Alternative Specifications

PPVT scores are measured for 3 and 4 year olds and PIAT performance for children two years older. One implication is that the samples differ (e.g. the PIAT assessments are conducted on children born to slightly younger mothers). To insure that the findings are not an artifact of these differences, I estimated models for the 3045 children with test scores reported for all three cognitive assessments. The results confirm more deleterious predicted effects of maternal employment on PIAT-R and PIAT-M achievement than on PPVT performance. An extra 20 hours per week of employment throughout the first three years is associated with -.077 (.034), -.108 (.036) and -.015 (.032) reductions in test scores (with standard errors in parentheses).

The impact of maternal job-holding was allowed to differ across high and low achievers through a series of quantile regression models examining the 10^{th}, 25^{th}, 50^{th}, 75^{th}, and 90^{th} percentiles of the test scores. These estimates failed to reveal any differential effects across achievement levels for the PIAT-R and PIAT-M assessments. However, employment was predicted to have more negative impacts for high than low PPVT scores: the coefficient (standard error) on average weekly hours in years 1 through 3 was .045 (.062), .004 (.040), -.044 (.029), -.063 (.030), and -.110 (.033) at the 10^{th}, 25^{th}, 50^{th}, 75^{th}, and 90^{th} percentiles.[57]

[57] Standard errors are bootstrapped estimates obtained by resampling the data 20 times.

Fixed-effect models were estimated by restricting the sample to siblings and including a vector of mother-specific dummy variables. As discussed, these models understate the costs of maternal employment if women supply less labor when their children have health or developmental problems. The results are consistent with this expectation (and the opposite of those anticipated if mothers with "problem" children return to work unusually rapidly). The FE estimates (standard errors) for a 20 hour per week increase in work hours throughout the first three years are .038 (.048), -.077 (.050) and -.067 (.052) for PPVT, PIAT-R and PIAT-M scores; corresponding OLS estimates are -.030 (.033), -.134 (.035), and -.097 (.035).[58] Thus, even with a likely upwards bias, the fixed-effect estimates suggest a negative relationship between early maternal employment and the reading or mathematics achievement of 5 and 6 year olds.

A final set of specifications controlled for the incidence or duration of breast-feeding, which has been linked to improved cognitive development (Anderson, Johnstone, and Remley, 1999). Working mothers are less likely to breast-feed, which could explain a portion of the negative effect of maternal employment (particularly during the first year). However, the results provide little support for this possibility. Breast-feeding is positively associated with the assessment scores but its inclusion only slightly reduces the magnitude of the year 1 labor supply coefficient and does not affect the parameter estimates for working in years 2 and 3.[59]

F. What About Fathers?

The preceding analysis suggests the importance of maternal investments in young children. But what about fathers? While it is possible that mothers provide unique inputs, it seems likely that there is at least some substitutability between parents. However, since men are typically paid more than women, larger income benefits could accrue to paternal employment. These issues are addressed in Table 6, which summarizes econometric results for children with both parents in the household on the interview date of the fourth calendar year after birth.

[58] Sample sizes are 3086, 3567, and 3656 for the three outcomes.

[59] The biggest change is that the year 1 coefficient in the PIAT-M equation increases from -.027 without controls for breast-feeding to -.025 (-.019) when the incidence (duration) of breast-feeding is held constant. Waldfogel et al. (forthcoming) similarly find that breastfeeding has insignificantly positive effects on cognitive scores but without affecting the estimate for maternal employment.

Parental employment refers to average weekly work hours (divided by 20) during years 1 through 3.[60]

Column (a), which holds constant maternal but not paternal hours, confirms that the labor supply of mothers has a small and insignificant negative predicted impact on the verbal ability of 3 and 4 year olds but substantial detrimental effects on the reading and math achievement of children aged 5 and 6 (although smaller sample sizes raise the standard errors).

Column (b) adds controls for paternal employment. This has virtually no effect on the maternal coefficients, indicating that the findings of previous sections are not biased by this omission, but suggests that the labor supply of fathers has either no impact or a *beneficial* consequence – working 20 extra hours per week is associated with .09, .03, and .04 standard deviation increases in PPVT, PIAT-R, and PIAT-M scores. This positive correlation is probably due to omitted variables bias. As discussed, joblessness is concentrated among a relatively small fraction of fathers likely to provide low quality investments. Regressions corresponding to specification (b), but controlling for weeks rather than hours worked, reveal an extremely strong relationship between weeks on the job and child test scores: switching from no work to labor supply in all 156 weeks is predicted to raise PPVT, PIAT-R, and PIAT-M performance by .38, .34, and .23 standard deviations (with standard errors of .10, .11, and .11). Such large positive correlations are expected if men employed few weeks have characteristics associated with poor child outcomes. Column (c) begins to address this possibility by holding constant the father's age and education (in the calendar year of the child's birth). Doing so reduces the coefficients on paternal employment by 8 to 12 percent.

The remainder of the table details the results of two strategies for better accounting for these sources of heterogeneity. Specification (d) averages paternal hours over weeks of work rather than all weeks.[61] Once this is done, there is never a large or significant positive correlation between paternal labor supply and child outcomes. Instead, a 20 hour per week increase in market work is associated with a .08 standard deviation reduction in PIAT-R scores, no change

[60] The number of missing observations was reduced by averaging work hours over the two years for which the data are available in cases where paternal hours are not reported in one of the three years.
[61] The less than one percent of children whose fathers do not work at all during the three years are excluded from this analysis.

in PIAT-M performance, and a modest .02 standard deviation rise in PPVT assessments. These "conditional" averages are useful if differences in weekly employment probabilities are unrelated to investments in children.

Columns (e) and (f) display coefficients for conditional work hours (of fathers) where the sample is restricted to children whose dads worked a minimum of either 117 or 143 weeks during the three years.[62] Given the high rates of paternal employment, relatively few children are excluded by these work restrictions (11-12 and 24-25 percent respectively). At a minimum, these specifications show the effects of paternal hours for the vast majority of children whose fathers hold jobs most of the time. More optimistically, they may control for the heterogeneity associated with differences in weeks worked. The coefficients on paternal employment are almost always negative in these models. For instance, in specification (f), an extra 20 hours per week of paternal employment reduces the anticipated math and reading performance of 5 to 6 year olds by .03 and .06 standard deviations. These decreases are 41 to 56 percent as large as those predicted for corresponding maternal labor supply. This suggests that time investments of mothers and fathers may have qualitatively similar effects and raises the possibility of substitution across parents. The stronger negative impacts observed for mothers could reflect actual differences. For instance, only women can breast-feed and men receive higher average wages implying potentially larger income effects. Alternatively, they may occur because the work hours and characteristics of males are less well measured.

G. Discussion

This analysis suggests that parental investments at the beginning of a child's life play a significant role in fostering cognitive development. Maternal employment in the first year is associated with lower verbal ability at ages 3 and 4, with partially offsetting benefits for labor supply during the next two years. Early job-holding has a more detrimental cumulative impact on the predicted reading and mathematics performance of 5 and 6 year olds, with negative effects persisting for work in the second and possibly the third year.

[62] Thus, model (e) requires the father to have worked in three-quarters of all weeks; column (f) implies that they were jobless for less than three months over the three years.

The coefficients on parental employment become consistently more negative (or less positive) with the addition of more complete controls for heterogeneity. This indicates that children with working parents come from relatively advantaged backgrounds or possess attributes associated with rapid cognitive development and makes it extremely unlikely that the observed detrimental effects result from remaining omitted variables bias. Instead, to the extent this heterogeneity has still not been fully accounted for, the estimates may continue to understate the deleterious impact of market work during the child's early years. A second implication is that much prior research, because it only crudely controls for these disparities, probably presents an overly optimistic assessment of these effects. This helps to explain why these studies have often predicted positive or neutral consequences of early job-holding, in contrast to the more deleterious impacts uncovered here. It also provides a reason why recent analyses, which typically account (somewhat) better for the heterogeneity, more uniformly indicate negative effects of labor supply by mothers during the child's first year.

Our confidence in these conclusions will be strengthened if future investigations confirm the findings and identify mechanisms for the effects. We need to better understand the role of child care and the impact of early employment on other child outcomes (such as health or socioemotional development). It is also important to explain why stronger negative effects are obtained for the PIAT scores of 5 and 6 year olds than for PPVT performance measured two years earlier. This does not result from changes in the sample composition but could occur because the tests measure different aspects of cognitive skill. Paternal labor supply is also particularly hard to examine, since the data are less adequate and joblessness is often unrelated to investments in young children. This analysis hints at the possibility of substitution between maternal and paternal investments but these results should be viewed as preliminary.

Notwithstanding these caveats, the estimates suggest cognitive costs to children whose mothers work excessive hours in the early years. An extra 20 hours per week of employment throughout the first three years is associated with .11 and .09 standard deviation reductions in the reading and math performance of 5 and 6 year olds. Losses of this size are qualitatively significant: they are equivalent to those predicted by a 2 to 3 year decrease in maternal education

and correspond to reductions in test scores from the median to the 46th and 47th percentiles. Currie & Thomas (1999) provide evidence that early test performance is strongly related to future educational and labor market outcomes, indicating that these effects may translate into lasting economic costs.[63]

Sixty-five (77) percent of the mothers in the representative portion of the NLSY return to work before their child's first (second) birthday. One possible reason is the concerted effort, since the mid-1980s, to increase the employment of women with young children through changes in welfare and Medicaid policies, the Earned Income Tax Credit, and government funding for child care (Meyer & Rosenbaum, 2001). A second is that rights to maternity leave are limited in the United States compared with other industrialized countries. Evidence of deleterious effects of maternal employment may therefore be relevant for these policies.

Parental investments during infancy are important. Not only is maternal job-holding associated with decreased verbal ability among 3 and 4 year olds, but work during the second and third years is correlated with relatively low performance on all three assessments if the mother also held a job during the first year. This suggests potential gains to children from expanded entitlements to parental leave or other methods of facilitating time at home with newborns.[64] The evidence further suggests that children do substantially better when their mothers work part-time during years 2 and 3 than if she is employed full-time. This indicates possible benefits of "family-friendly" policies that promote a gradual return to full involvement in the labor market.

Child cognitive development is just one argument in the household utility function, however, raising the possibility of tradeoffs between this and other desirable outcomes. Moreover, the benefits of early parental investments might be partially or fully offset by reductions in future incomes if the time away from work adversely affects labor market advancement. The consequences of parental employment are also likely to depend on the technologies and institutional arrangements in place. For example, many European countries

[63] Currie & Thomas analyze the effects of reading test scores at age 7 on outcomes at age 33, using data from the British National Child Development Survey.
[64] Ruhm (2000) provides evidence that parental leave entitlements also improve child health.

heavily subsidize child care as a first step in the system of public education. The effects of early job-holding could be quite different in such an environment. Finally, better understanding the mechanisms by which parental investments promote child cognitive development might facilitate designing less costly ways of achieving the same benefits. These represent important topics for future research.

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Appendix

Table A.1: Variables Used in Analysis

Variable	Description
	Outcomes
PPVT	Peabody Picture Vocabulary Test-Revised: 3-4 year Olds
PIAT-M	Peabody Individual. Achievement Test, Mathematics: 5-6 year olds
PIAT-R	Peabody Individual. Achievement Test, Reading Recognition: 5-6 year olds
	Maternal Employment
Hours Worked	Hours worked (divided by 20) during specified period
Weeks Worked	Proportion of weeks worked (divided by 2) during specified period
Weeks Resume	Weeks after birth of child until mother resumes employment
	"Basic" Child, Maternal, and Household Characteristics (B)
Age	Age of child (in months) at assessment date
Age Squared	Age Squared of child at assessment date
Race/Ethnicity	Child is Hispanic or a non-Hispanic Black (2 d.v.'s)
Female	Child is Female (d.v.)
Parity	Birth order of child
Residence	Lives in SMSA or central city at assessment date (2 d.v.'s)
Mother's Age	Age (in years) of mother at the time of child's birth
Education	Highest grade completed by mother at assessment date
	Supplemental Maternal, Family, and Child Characteristics (S)
Father Present	Father living in household at assessment date (d.v.)
Married	Mother is married at survey date before pregnancy and next 3 surveys (4 d.v.'s)
Divorced	Mother is separated, divorced, or widowed at survey date before pregnancy and next 3 surveys (4 d.v.'s)
Poverty	Family income below poverty line in 2nd calendar year before child's birth (d.v.)
Siblings	Mother's number of siblings (top-coded at 15)
AFQT Score	Mother's score on the Armed Forces Qualification Test in 1980
Foreign Born	Mother born outside the United States (d.v.)
Foreign Language	Foreign language spoken in the mother's home at age 14 (d.v.)
Location	Mother lived outside U.S., in Southern U.S., or in rural area at age 14 (3 d.v.'s)
Magazines	Mother had magazines in home at age 14 (d.v.)
Newspaper	Mother's family received a newspaper at age 14 (d.v.)
Library Card	Someone in mother's household had a library card at age 14 (d.v.)
Grandmother1	Mother's mother was born outside the United States (d.v.)
Grandmother2	Highest grade completed by mother's mother
Grandfather1	Mother's father was born outside the United States (d.v.)
Grandfather2	Highest grade completed by mother's father
Private	Mother's current or last secondary school attended in 1979 was private (d.v.)
Both Parents	Mother lived with both mother and father at age 14 (d.v.)
Mother Only	Mother lived with mother but no male in household at age 14 (d.v.)
Siblings	Sibling born $\leq$18, 19-36 months before/after child's birth (4 d.v.'s)
Religion	Mother's religion was Baptist, Catholic, or no affiliation in 1979 (3 d.v.'s)
Church	Mother attended church < once, once, > once per week in 1979 (3 d.v.'s)
First Smoked	Mother smoked first cigarette by $\leq$10, 11-13, 14-16 years old, had not smoked by 1984 (4 d.v.'s)
Drug Use	Mother had tried marijuana/hashish, amphetamines/stimulants, cocaine, other drugs (barbiturates, sedatives, tranquilizers, psychedelics, heroin, other narcotics, or inhalants) by age 21 (4 d.v.'s)
Attitudes	Mother's cumulative Score on 8 Family Roles Attitude Questions

Table A.1 (Continued)

Gestation	Length of Gestation <35, 35, or 36-37 weeks (3 d.v.'s)
Birth Weight	Birth weight <57, 57-72, 73-88, or 89-112 ounces (4 d.v.'s)
Long Hospital Stay	Child stayed in hospital longer than mother following birth (d.v.)
Hospitalization	Child was hospitalized during first year of life (d.v.)
Doctor Visit	Child visited doctor due to illness during first year of life (d.v.)
Prenatal Care	1st Prenatal Care in 1-2, 3-4, 5-6 month of pregnancy (3 d.v.'s)
Alcohol	Mother drank at least one alcoholic beverage per month during pregnancy (d.v.)
Smoker	Mother smoked during pregnancy (d.v.)
Weight Gain	Weight gain during pregnancy <0, 0-15, >50 lbs. (3 d.v.'s.)
Body Mass Index	Mother's BMI before pregnancy <18.5, 25-30, 30-35, >35 (4 d.v.'s)
	Maternal Employment Characteristics (E)
Hours	Average weekly work hours (divided by 20) in Year 4
Wages	Hourly wages ($1996) in 4th quarter prior birth were: <$5, $7-$7.49, $7.50-$10, $10-$15, $15-$20, >$20 (6 d.v.'s)
Occupation	Occupation of main job in 4th quarter prior to birth was: professional/technical, managerial, sales, clerical, operative, service (6 d.v.'s)
	Other Regressors
Income	Family income ($1996) in calendar year before birth and next 4 years (5 variables)
Day Care	Child in regular nonparental care in specified year after birth
Day Care Mode	First day care arrangement during specified period provided by relative, nonrelative, or in group care center/nursery school/preschool (3 d.v.'s)
	Paternal Variables
Paternal Hours	Father's average weekly work hours in years 1 through 3. Year 1 refers to the calendar year of birth (the next year) for children born before (on or after) July 1
Paternal Age	Father's age in calendar year of child's birth
Paternal Education	Father's highest grade completed in calendar year of child's birth

Note: All variables are obtained from the NLSY. Body Mass Index is calculated using weight immediately before pregnancy and height at the 1981 interview.

Table 1: Previous Research Examining The Effects of Parental Employment On Cognitive Development			
Study	**Data/ Sample**	**Results**	**Comments**
Baydar & Brooks-Gunn (1991)	NLSY, 572 white 3-4 year olds (in 1986)	Maternal employment during the first year of life has negative effects on PPVT. No negative effects for working in second or third year. Possible positive effects for working all three years. Some evidence of bigger negative effects for entering work earlier in the first year.	Other controls limited to maternal AFQT score, child gender, parity, poverty status. Nonlinear relationship between work hours in first year and outcomes.
Blau & Grossberg (1992)	NLSY, 874 3-4 year olds (in 1986)	Maternal employment in the first (second & third) years of life associated with lower (higher) PPVT scores. No net effect of working in all three years. Much of second & third year benefit is due to higher incomes. Larger negative first year effects for high income households.	Controls for parent's education, household incomes, race/ethnicity, gender, parity, % of life in female-headed household. IV models estimated, but few plausible instruments.
Brooks-Gunn, Han, & Waldfogel (forthcoming)	NICHD, 900 non-Hispanic whites born in 1991	Employment in first 9 months of child's life negatively associated with Bracken School Readiness scores at age 3. Bigger effect for full-time work (>30 hours per week), low maternal sensitivity, boys, and married women. Smaller effects on MDI scores at 15 or 24 months of age.	Controls for home environment, maternal sensitivity, child care mode and quality, maternal characteristics. Pattern of employment effects suggests presence of selection bias.
Desai, Chase-Lansdale, & Michael (1989)	NLSY, 503 4 year olds (in 1986)	Negative effect of maternal employment on PPVT scores, particularly for continuous employment in first 4 years. Stronger negative effect of maternal job-holding in the first year for boys in high income families. Possible positive effect of employment beginning in the second year for girls.	Controls for maternal characteristics (age, verbal ability, education, marital history, race/ethnicity), household income, birth order, sibling age, number of day care arrangements.
Ermisch & Francesconi (2001)	BHPS, 591 siblings born in 1970-1979	Negative effect of full-time maternal employment during first five years of life on probability of subsequently completing A level exams in sibling fixed-effect models. No effect of part-time jobs or paternal employment.	Controls for age, gender, parity, household structure, maternal and paternal characteristics. Small sample sizes and parental labor supply crudely measured.
Greenstein (1995)	NLSY, 2040 4-6 year olds (in 1986, 1988, 1990)	Little relationship between maternal employment and PPVT score. Interactions between maternal employment and family income or cognitive stimulation also insignificant. No evidence that maternal employment has more detrimental effects for high income households.	Controls for child and maternal characteristics and family environment. Results difficult to interpret because of potential collinearity between regressors.
Han, Waldfogel, and Brooks-Gunn (2001)	NLSY, 412 black and whites born in 1982 or 1983	Negative effects of first year employment on PPVT scores of 3-4 year olds and PIAT scores of 5-8 year olds. Larger negative impact of full-time than part-time work but some offsetting benefits of working in years 2 and 3.	Controls for sex, parity, family poverty status, and maternal characteristics and household characteristics. Small sample sizes reduce statistical power.
Harvey (1999)	NLSY, 3-12 year olds in 1986, 1988, 1990, 1992, 1994	Negative effect of maternal work hours on PPVT and PIAT scores at young ages, which weakens or disappears at later ages. Some differences with marital status, income, or paternal employment.	Controls for individual, maternal, household characteristics. Regressors and samples vary across models, so results difficult to interpret, often subject to omitted variables bias.
Hill et al., (2001)	NLSY, 4143 Children born from 1982-1989	Negative effects of maternal employment in first year (particularly full-time work) on PPVT scores of 3-4 year olds and PIAT scores of 5-8 year olds. No impact of labor supply in second or third year. Few differences by race or sex; some evidence of stronger negative effects for children in married or high income households.	Propensity score matching methods used to account for heterogeneity. One consequence is that sample sizes are often small. The dependent variable is also sometimes imputed rather than observed.

James-Burdumy (1999)	NLSY, 2119 3-4 year olds in 1986 or 1988	OLS estimates indicate negative (positive) effects of maternal employment during first (second) year on PPVT scores; however, no consistent employment effect was obtained for any of the first three years in models that include maternal fixed-effects.	Controls for individual, maternal and household characteristics. FE and IV techniques used to control for heterogeneity but with small samples (498 siblings) and weak instruments.
Leibowitz (1977)	Sesame data, 805 3-5 year olds (in 1969)	Full-time maternal employment negatively (but not quite significantly) associated with PPVT. No effect of part-time employment. Positive effect of labor saving devices (dishwashers) in home and of reading to children/self, but negative effects of other activities. Negative relationship between number of children & PPVT.	Controls for parent's education, race/ethnicity, native language, home environment, day care arrangements, number of children.
Moore & Driscoll (1997)	NLSY, 1154 5-14 year olds (1992); mothers on AFDC in 1986-1990	Maternal employment in 1991 associated with higher PIAT Reading and Math scores. Most effects eliminated after controlling for child, maternal, & household characteristics, although behavioral problems and higher math scores persist for daughters of higher earning women.	Controls for sex, age, birth order, health, birth weight, maternal characteristics & attitudes, family employment & AFDC history. Omitted variable bias probably persists and may explain benefits of working.
Mott (1991)	NLSY, 2387 1-4 year olds (in 1986)	Maternal employment averaging more than 20 hours per week in second quarter of child's life negatively related to PPVT scores (ages 3-4); no effect on MFL scores (ages 1-3). No effect of lower work hours. Employment in first quarter insignificantly negatively related to MFL scores.	Comprehensive controls for child & family characteristics including early health problems and substance use during pregnancy. Collinearity between maternal employment and child care arrangements makes results difficult to interpret.
Neidell, 2000	NLSY, 1681-4581 children (age and survey date not specified)	Delaying resuming maternal employment until the 6^{th} through 12^{th} month of child's life has a marginally significant (insignificant) positive effect on PPVT (PIAT) scores. Ambiguous impacts of returning to jobs at late ages. Models that less adequately control for heterogeneity imply harmful effects of early employment.	Controls for child age, birth-weight, breastfeeding, maternal/ family characteristics, and sibling fixed-effect. Maternal employment limited to dichotomous measures of weeks after birth until return to work.
Parcel & Menaghan (1994)	NLSY, 768 3-6 year olds with employed mothers (in 1986)	Maternal employment during the first year or first three years positively correlated with PPVT scores. Effect is not monotonic in work hours. Subsequent full-time maternal work correlated with higher PPVT, compared to working part-time. Early (current) full-time *paternal* employment associated with insignificantly lower (higher) PPVT scores. Some evidence that maternal employment is more problematic for less well-off women.	Comprehensive controls for child and parent characteristics, the home environment, and working conditions. Results are difficult to interpret because of potential endogeneity (e.g. home environment and work hours) and collinearity (e.g. wage, occupational complexity, and work hours) of regressors.
Stafford (1987)	SRC Time Use Study, 77 elem. school students (in 1981/2)	Cognitive skills (measured by 7 indicators and a composite) fall with the number of siblings (particularly males), rise with family income, and decline with mother's market work hours in the pre-school years.	Detailed teacher evaluations of cognitive development. Maternal employment refers to various ages prior to start of school.
Vandell & Ramanan (1992)	NLSY, 189 low income non-Hispanic second-graders (in 1986)	Maternal employment in first three years (and subsequently) positively correlated with the mothers' education, AFQT score, family income, and quality of home environment. Employment in first three years correlated with higher PIAT-math and insignificantly lower PPVT scores. Recent employment correlated with higher PIAT-reading and PPVT scores.	Controls for the child's race, gender; mother's age, education, marital status, family income, attitudes, AFQT and Rosenberg self-esteem scores, and HOME scale. Sample selection criteria is not specified.

Waldfogel, Han & Brooks-Gunn (forth-coming)	NLSY, 1872 children born in 1982-89, assessed at ages 3-8	Negative effects of maternal employment in the first year on PPVT and PIAT scores for white or Hispanic children (particularly for girls), which increase with the number of hours worked. Ambiguous results for work in the second or third year and little impact on African-American children. Smaller deleterious employment effects in sibling fixed-effect models	Controls for sex, parity, family income, maternal characteristics, breast feeding, day care arrangements. Maternal employment mostly measured by dichotomous variables for any work versus no work.

Notes: Abbreviations: NLSY: National Longitudinal Survey of Youth; SRC: Survey Research Center; NICHD: NICHD Study of Early Child Care; BHPS: British Household Panel Survey; MDI: Bayley Mental Development Index; MFL: Memory for Location score; PIAT: Peabody Individual Achievement Test; PPVT: Peabody Picture Vocabulary Test.

Table 2: Sample Means of Selected Variables By Average Weekly Work Hours of Mother in Specified Periods

Variable	Full Sample	Work Hours During Year 1			Work Hours During Years 2 and 3		
		0	1-29	≥30	0	1-29	≥30
Cognitive Outcomes							
PPVT	0.00	-0.15	0.09	0.10	-0.21	0.08	0.07
	(.02)	(.04)	(.03)	(.04)	(.05)	(.03)	(.03)
PIAT-R	0.00	-0.10	0.04	0.11	-0.17	0.03	0.10
	(.02)	(.03)	(.03)	(.04)	(.04)	(.03)	(.04)
PIAT-M	0.00	-0.11	0.05	0.11	-0.16	0.04	0.08
	(.02)	(.03)	(.03)	(.04)	(.04)	(.03)	(.03)
Family Background							
Mother's Age at Birth	24.7	24.3	24.4	25.9	24.8	24.1	25.6
(years)	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)
Mother's Education	12.5	12.0	12.7	13.2	12.0	12.4	13.1
(years)	(.04)	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)	(0.1)
Mother's AFQT Score	67.9	61.9	70.6	73.8	61.5	68.4	72.6
	(0.4)	(0.6)	(0.6)	(0.7)	(0.8)	(0.6)	(0.6)
Mother Married at Birth (%)	62.9	57.6	64.1	70.3	59.5	62.8	65.8
	(0.9)	(1.4)	(1.4)	(1.7)	(1.7)	(1.3)	(1.6)
Mother's Work Hours in 4th	21.7	11.3	23.4	37.4	10.7	20.3	33.7
Quarter Before Birth	(0.3)	(0.5)	(0.5)	(0.5)	(0.6)	(0.5)	(0.5)
Father in Household in 4th	75.8	71.2	77.9	80.3	72.5	76.9	76.5
calendar year after birth (%)	(0.8)	(1.3)	(1.2)	(1.5)	(1.6)	(1.1)	(1.4)
Family in Poverty in 2nd	18.8	29.9	14.4	6.3	31.2	18.5	8.4
Year Before Birth (%)	(0.7)	(1.4)	(1.1)	(0.9)	(1.7)	(1.1)	(0.1)
Child Characteristics							
Low Birth Weight (%)	6.5	7.7	5.8	5.5	8.6	5.4	6.2
	(0.4)	(0.8)	(0.7)	(0.9)	(1.0)	(0.6)	(0.8)
Preterm Birth (%)	19.7	19.6	19.2	20.9	21.0	18.0	21.1
	(0.7)	(1.1)	(1.2)	(1.5)	(1.4)	(1.0)	(1.4)
Hospitalized During 1st	7.4	8.4	7.3	5.7	7.8	8.2	5.9
Year of Life (%)	(0.5)	(0.8)	(0.8)	(0.9)	(0.9)	(0.7)	(0.8)

Notes: Table displays averages for the nationally representative subsample of the NLSY The size of the full sample is 3,042; missing information reduces the sample size for some variables. Year 1 refers to the first 4 quarters after birth; years 2 and 3 to the fifth through twelfth quarters after birth. Standard errors of sample means are in parentheses. The PPVT, PIAT-R and PIAT-M indicate standard scores, normalized to have a standard deviation of one, on the Peabody Picture Vocabulary Test-Revised, and the Peabody Individual Achievement Reading Recognition and Mathematics Subtests. The sample for these tests are 5-6 year (60-83 month) old children, except for the PPVT, where 3-4 year (36-59 month) old children are assessed. In most cases, information on PPVT and PIAT scores are obtained for the same child two years apart. Mother's education is measured at the time of the PIAT assessments. Low birth weight indicates that the child weighed less than 5.5 pounds at birth. Preterm birth refers to gestation of less than 38 weeks.

Table 3: Regression Estimates of the Effects of Maternal Employment on Child Cognitive Development

Time Period	(a)	(b)	(c)	(d)	(e)	(f)	(g)	(h)	(i)
					PPVT Score				
Year 1	.054	-.039	-.038	-.044	-.065	-.064	-.090	-.066	-.052
	(.029)	(.024)	(.023)	(.024)	(.027)	(.027)	(.027)	(.024)	(.026)
Years 2 and 3	.097	.056	.023	.030	.030	.027	.022	.019	.024
	(.027)	(.023)	(.023)	(.028)	(.028)	(.028)	(.028)	(.024)	(.028)
Year 0					.050	.046	.085	.066	.041
					(.028)	(.028)	(.027)	(.025)	(.031)
					PIAT-R Score				
Year 1	.091	.020	.012	-.015	-.028	-.030	-.028	-.029	-.056
	(.027)	(.025)	(.024)	(.026)	(.029)	(.029)	(.028)	(.025)	(.027)
Years 2 and 3	.055	-.010	-.045	-.081	-.081	-.080	-.076	-.068	-.061
	(.026)	(.024)	(.024)	(.029)	(.029)	(.029)	(.029)	(.026)	(.029)
Year 0					.030	.029	.033	.017	.068
					(.029)	(.029)	(.029)	(.026)	(.032)
					PIAT-M Score				
Year 1	.082	.017	.007	-.011	-.027	-.030	-.030	-.031	-.025
	(.027)	(.025)	(.025)	(.026)	(.029)	(.029)	(.029)	(.026)	(.028)
Years 2 and 3	.041	-3.2E-4	-.033	-.058	-.059	-.058	-.062	-.048	-.063
	(.026)	(.024)	(.024)	(.030)	(.030)	(.030)	(.030)	(.026)	(.030)
Year 0					.037	.035	.043	.018	.061
					(.030)	(.030)	(.030)	(.026)	(.033)
Additional Regressors	None	B	B,S	B,S,E	B,S,E	B,S,E,I	B,S,E	B,S,E	B,S,E
Dependent Variable	Standard Score	Standard Score	Standard Score	Standard Score	Standard Score	Standard Score	Percentile Score	Raw Score	Standard Score
Employment Variable	Hours	Hours	Hours	Hours	Hours	Hours	Hours	Hours	Weeks

Notes: Table shows coefficients for OLS regressions of the specified assessment score, normalized to have a standard deviation of 1. Standard errors are shown in parentheses. PPVT scores are measured for children 36-59 months of age; PIAT scores for those aged 60-83 months. "Standard" scores are used in all specifications except columns (g) and (h), where percentile and raw scores are used. "Hours" indicate average weekly work hours divided by 20. "Weeks" refers to the proportion of weeks worked multiplied by two during the specified period. Year 1 includes the first four quarters after birth; Years 2 and 3 refer to the fifth through twelfth quarters after birth, year 4 to the thirteenth through sixteenth quarters subsequent to it, and Year 0 to the four quarters *before* birth. All models control for the assessment year. The categories of additional regressors are "Basic" Child, Maternal, and Household Characteristics (B); Supplemental Maternal, Family, and Child Characteristics (S); Maternal Employment Characteristics (E), and Family Incomes (I) in the calendar year before birth and the next 4 years. Full descriptions are provided in Table A.1. Sample sizes are 4,180, 4,695, and 4,803 for the analysis of PPVT, PIAT-R, and PIAT-M scores. Larger sample sizes for the PIAT scores primarily reflect inclusion of the relatively large number of children born in 1980 and 1981, who were 5 or 6 years old in 1986 (the first assessment year) and for whom PPVT scores were not obtained at 3 or 4 years of age.

Table 4: Alternative Regression Estimates of the Effects of Maternal Employment

Employment Regressor	PPVT	PIAT-R	PIAT-M
Any Work in Year 1	-.046	-.080	-.086
	(.034)	(.035)	(.036)
Any Work in Years 2 or 3	.091	.039	.068
	(.036)	(.038)	(.038)
Year 1: Part-Time vs. No Work	-.037	-.063	-.072
	(.034)	(.036)	(.037)
Year 1: Full-Time vs. Part-Time	-.038	-.040	.004
	(.041)	(.045)	(.045)
Years 2 and 3: Part-Time vs. No Work	.081	.035	.065
	(.036)	(.038)	(.038)
Years 2 and 3: Full-Time vs. Part-Time	-.041	-.091	-.109
	(.040)	(.042)	(.043)
Resumes Work: 27-52 Weeks After Birth	.023	.022	-.042
	(.045)	(.046)	(.047)
Resumes Work: 53-104 Weeks After Birth	.120	.128	.088
	(.044)	(.045)	(.046)
Resumes Work: 105-156 Weeks After Birth	-.003	.127	.083
	(.054)	(.057)	(.059)
Resumes Work: >156 Weeks After Birth	.052	.066	.045
	(.043)	(.044)	(.045)
Work Hours in Year 1	-.060	-.017	6.6E-4
	(.031)	(.033)	(.033)
Hours in Years 2 and 3 if Hours in Year 1 = 0	.040	-.058	6.9E-4
	(.042)	(.045)	(.046)
Hours in Years 2 and 3 if Hours in Year 1 > 0	.026	-.089	-.079
	(.042)	(.032)	(.032)

Notes: See note on Table 3. The top panel includes dichotomous variables indicating if the mother worked any hours during the specified period. The second panel estimates the effects of part-time and full-time work (vs. no employment), with 30 hours per week as the dividing line between part-time and full-time jobs. The third panels controls for the number of weeks after birth until the mother resumes employment, with 0-26 weeks the reference group. Work hours during year 4 are not controlled for in this panel. The last panel allows the effects of maternal employment in years 2 and 3 to differ depending on whether or not the mother worked in year 1. Other than these changes, the control are the same as in specification (e) of Table 3.

Table 5: Regression Estimates of the Effect of Maternal Employment on Alternative Groups of Children

Population Group	PPVT		PIAT-R		PIAT-M	
	Years 1 - 3	Year 0	Years 1 - 3	Year 0	Years 1 - 3	Year 0
Full Sample	-.025 (.028)	.030 (.026)	-.111 (.030)	.033 (.027)	-.086 (.030)	.037 (.028)
Boys	-.027 (.042)	.058 (.040)	-.080 (.042)	.049 (.041)	-.082 (.044)	.079 (.043)
Girls	-.009 (.038)	-.003 (.034)	-.128 (.043)	.006 (.038)	-.068 (.043)	.008 (.043)
Whites	-.033 (.035)	.039 (.033)	-.090 (.042)	.012 (.038)	-.047 (.042)	.013 (.039)
Blacks	.042 (.056)	-.034 (.051)	-.101 (.058)	-.006 (.058)	-.065 (.061)	.052 (.056)
Spouse Present	-.049 (.036)	.023 (.033)	-.095 (.039)	-.004 (.036)	-.056 (.040)	.002 (.036)
Spouse Not Present	.127 (.062)	.068 (.058)	-.021 (.062)	.045 (.060)	.038 (.069)	.038 (.066)
Wage ≥ $10/Hour	-.091 (.063)	.034 (.061)	-.081 (.070)	-.029 (.069)	-.030 (.071)	-.019 (.069)
Wage < $10/Hour	.010 (.037)	.038 (.032)	-.102 (.042)	.040 (.036)	-.068 (.043)	.048 (.036)

Notes: See notes on Table 3. This table shows coefficients on average weekly work hours during the year before birth and the first three years of the child's life. These are obtained from models controlling for the same variables as specification (e) of Table 3. The second panel divides the sample between boys and girls, while the third provides separate estimates for non-Hispanic whites and non-Hispanic blacks (with Hispanics excluded). The sample in the first row of the fourth panel is restricted to children whose mothers have a spouse in the household at each of the survey dates in the first three calendar years following birth. The second row refers to children whose mothers do not have a spouse living in the household at any of the three survey dates. The bottom panel stratifies the sample by the hourly wage (in $1996) of the mother in the fourth quarter prior to the child's birth; this analysis is restricted to women reporting wages at that time.

Table 6: Estimated Effects of Maternal and Paternal Employment During the First Three Years

Type of Parental Employment	(a)	(b)	(c)	(d)	(e)	(f)
			PPVT Score			
Mother	-.025 (.036)	-.023 (.036)	-.021 (.036)	-.023 (.036)	-.012 (.037)	-.035 (.040)
Father		.086 (.034)	.079 (.034)	.016 (.043)	-.002 (.046)	.002 (.051)
			PIAT-R Score			
Mother	-.077 (.041)	-.077 (.041)	-.069 (.041)	-.070 (.041)	-.094 (.044)	-.111 (.048)
Father		.030 (.039)	.026 (.039)	-.073 (.049)	-.057 (.055)	-.062 (.061)
			PIAT-M Score			
Mother	-.074 (.041)	-.073 (.041)	-.067 (.041)	-.067 (.041)	-.079 (.044)	-.080 (.048)
Father		.042 (.039)	.037 (.039)	-.002 (.049)	-.007 (.055)	-.033 (.061)
Paternal Characteristics	No	No	Yes	Yes	Yes	Yes
Averaging of Paternal Hours	N/A	All Weeks	All Weeks	Weeks Worked	Weeks Worked	Weeks Worked
Minimum # of Weeks Worked By Father	No Restriction	No Restriction	No Restriction	No Restriction	≥ 117	≥ 143

Notes: See note on Table 3. The models control for the same regressors as specification (e) of that table, except that maternal employment is averaged over the first three years. Columns (b) through (f) also include the father's average weekly work hours (divided by 20) in the first three years, as well as paternal hours in years 0 and 4. The sample is limited to children whose father resided in the mother's household on the survey date of the fourth calendar year following the birth and for whom information on paternal employment is available in all three years. Columns (c) through (f) also control for the father's age and years of schooling in the calendar year of the child's birth. Fathers' work hours during the first three years are averaged over all weeks in specifications (b) and (c) and over weeks of employment only in models (d) through (f). In specification (d), these "conditional" estimates exclude children whose fathers were never employed during the first three years. Models (e) and (f) further restrict the sample to children whose fathers worked at least 117 and 143 weeks during the three years. Sample sizes are 2383, 2517, and 2579 for PPVT, PIAT-R, and PIAT-M scores for the full sample. For the corresponding conditional estimates they are 2367, 2499, and 2561. Sizes of the subsample with fathers working at least 117 weeks are 2099, 2198, and 2251, and for those with fathers employed 143 weeks or more they are 1805, 1884, and 1930.

Fig. 1: Weeks Worked By Mothers Before and After Child Birth

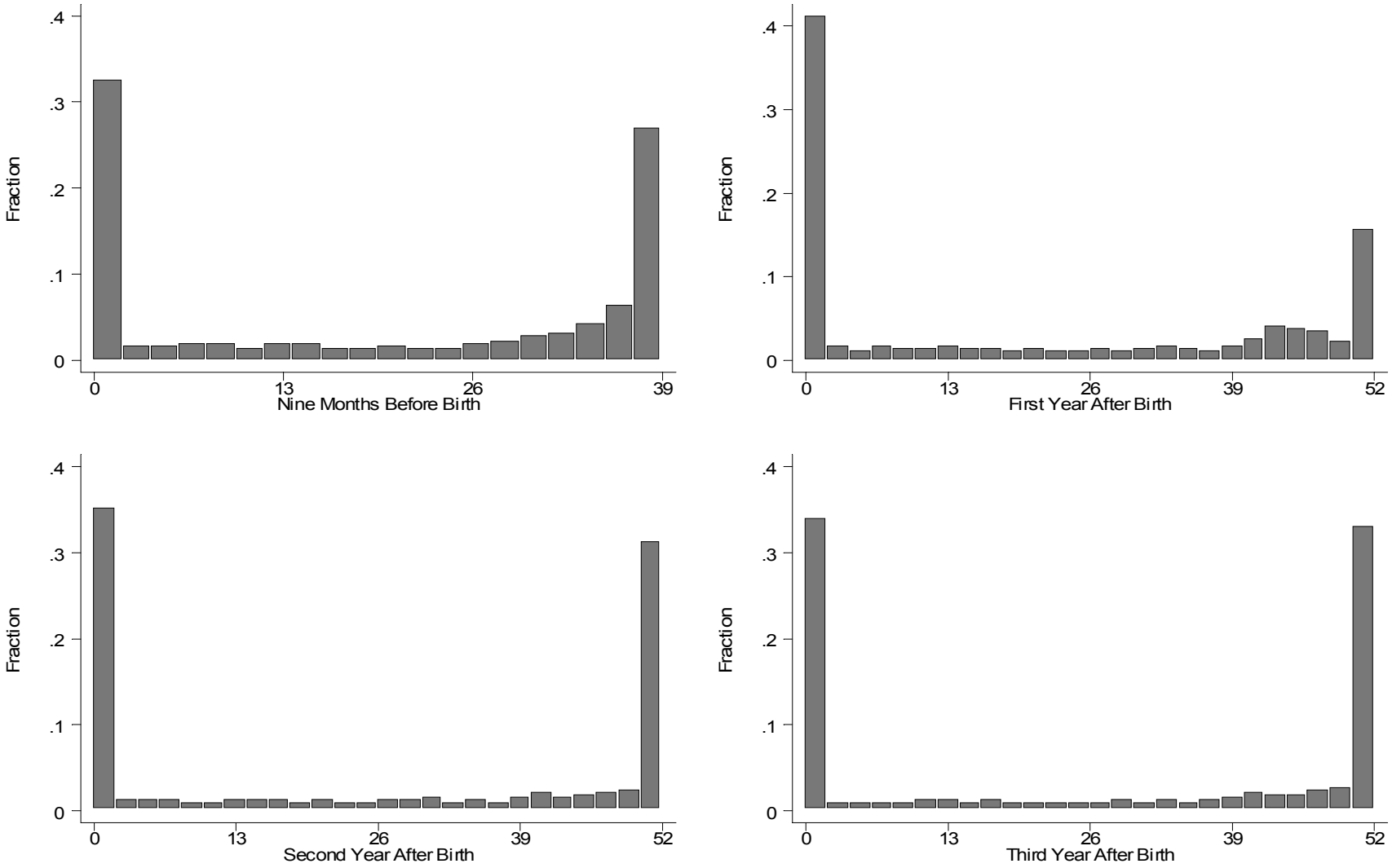

Fig 2: Weekly Work Hours of Mothers Before and After Child Birth

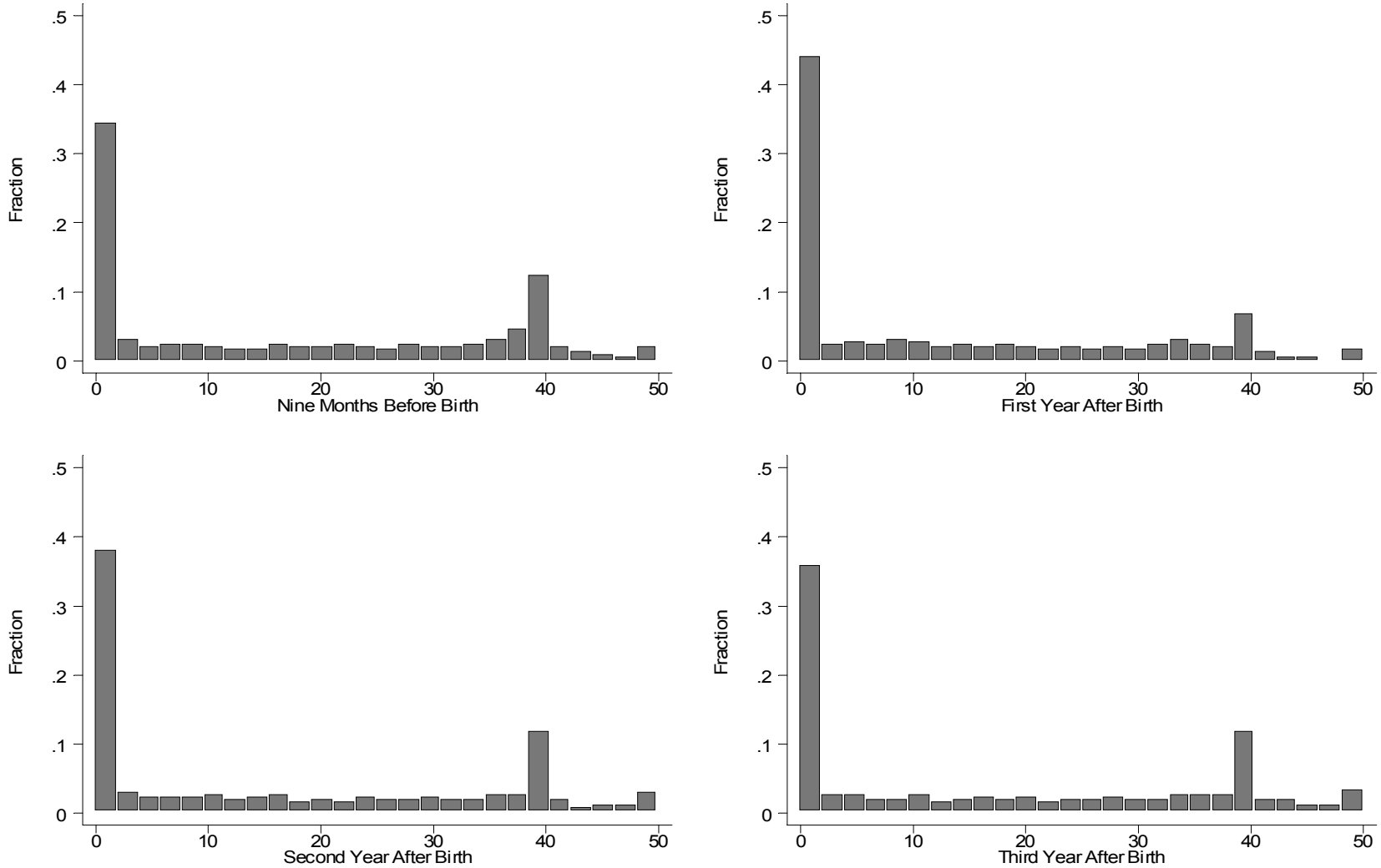

Fig 3: Average Work Hours Of Mothers In Weeks of Employment

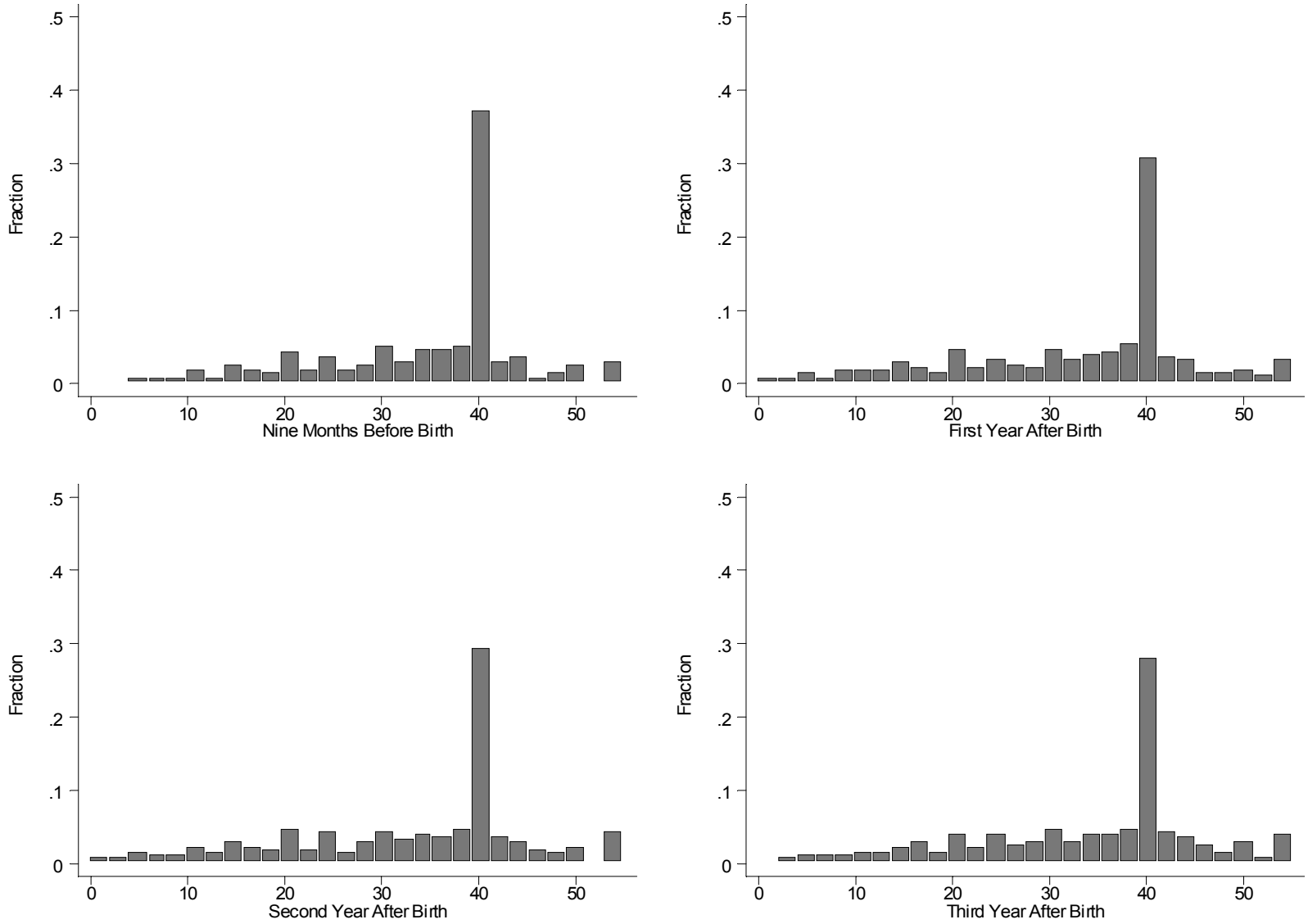

Fig. 4: Employment and Hazard Rates of Pregnant Women In Weeks Before Birth

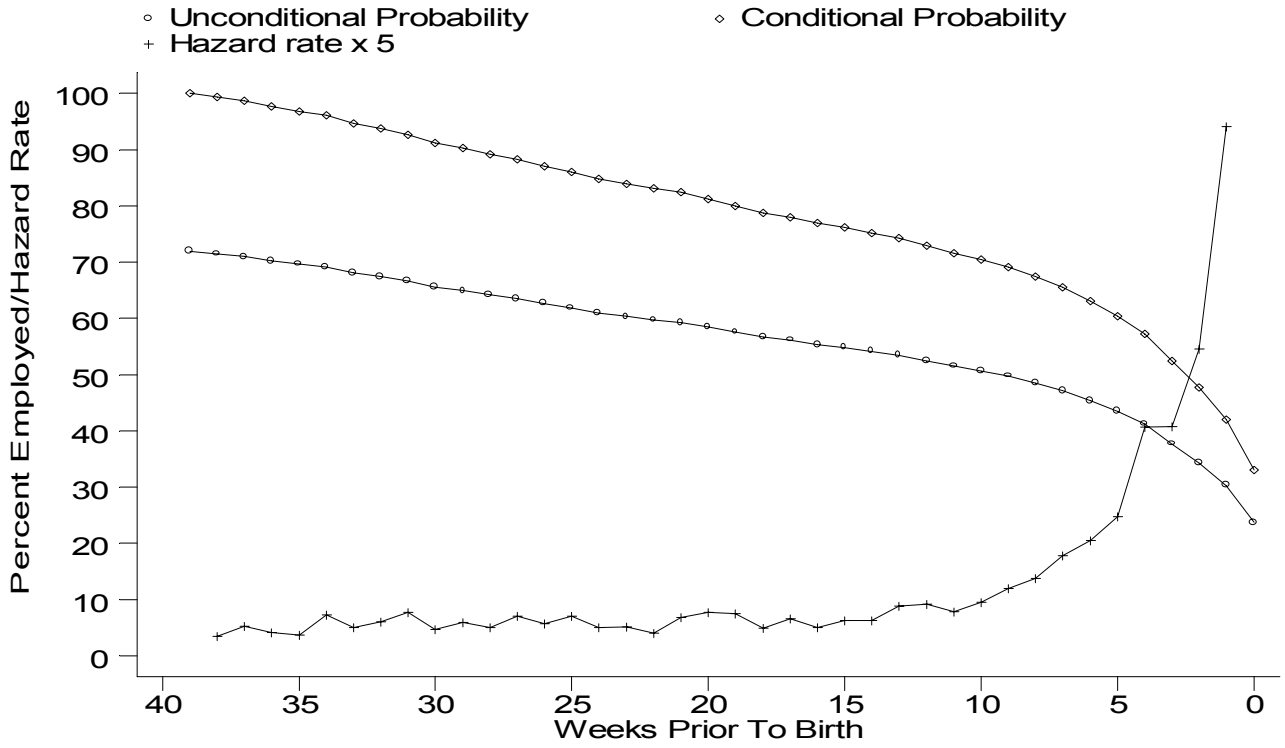

Fig. 5: Reemployment Probabilities and Hazard Rates of Mothers By Child Age (in weeks)

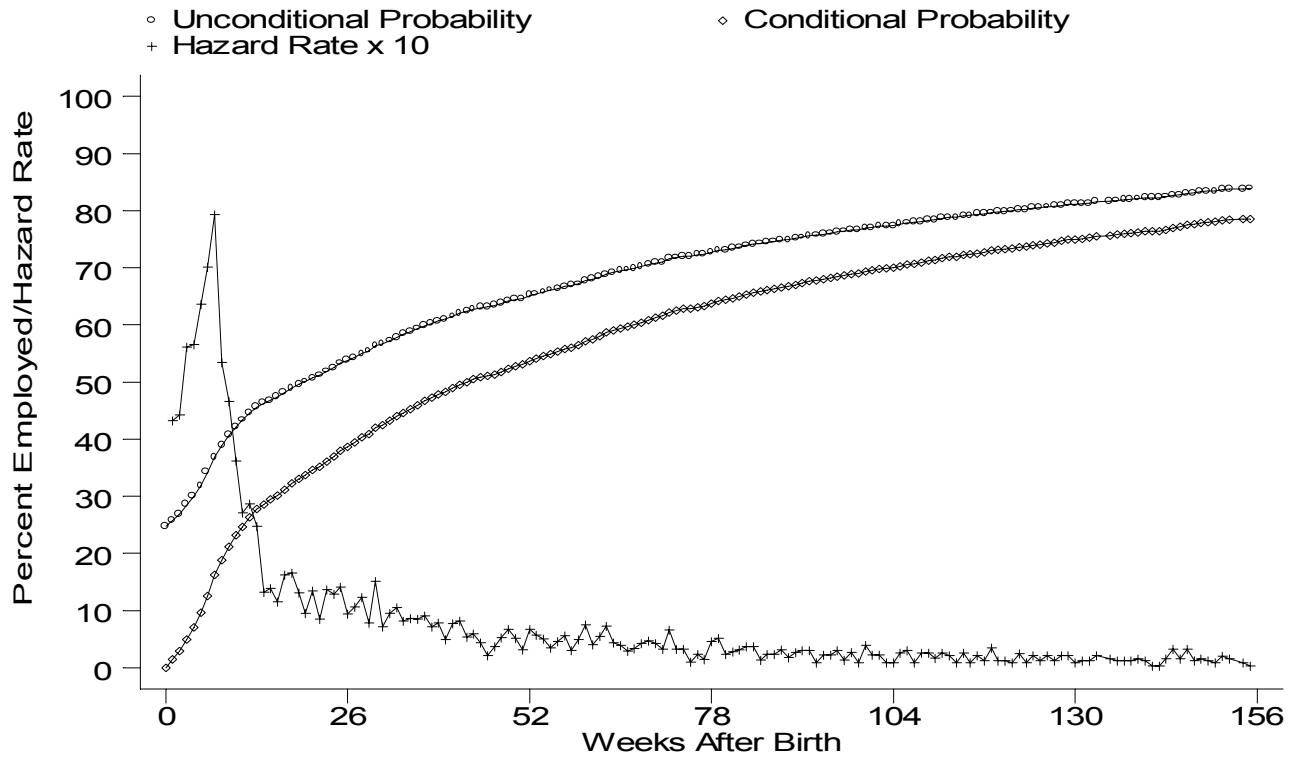

Fig. 6: Weeks and Hours Worked By Fathers