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DOES RETURNING TO WORK AFTER CHILDBIRTH AFFECT BREASTFEEDING PRACTICES?

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ABSTRACT

Although the Surgeon General recently highlighted breastfeeding as "......one of the most important contributors to infant health," few health economics studies based in developed countries have considered breastfeeding as an important health behavior that can be influenced by labor market decisions and by public policies. This study examines the effect of the timing and intensity of returning to work after childbirth on the probability of initiating breastfeeding and the number of weeks of breastfeeding. Data come from the National Longitudinal Survey of Youth (NLSY79). Baseline probit models and family-level fixed effects models indicate that returning to work within 3 months is associated with a reduction in the probability that the mother will initiate breastfeeding by 16-18%. Among those mothers who initiate breastfeeding, returning to work within 3 months is associated with a reduction in the length of breastfeeding of 4-6 weeks. We find less consistent evidence that working at least 35 hours per week (among mothers who return to work within 3 months) detracts from breastfeeding. Baseline and fixed effects models indicate that returning to full-time work is associated with a reduction in the length of breastfeeding of 1-4 weeks; however, we do not find consistent evidence regarding the association between returning to full-time work and breastfeeding initiation. Overall, the findings suggest that maternal employment is negatively associated with both breastfeeding initiation and breastfeeding duration.

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1. INTRODUCTION

As of the year 2000, 51% of mothers of infants less than 12 months old worked outside of the home [1]. The number of employed mothers with infants has grown dramatically since 1976, when the government first started to collect these data. Many researchers have studied the effects of maternal employment on child development. Recent studies suggest that some forms of maternal employment during the child's first year may have negative effects on children's cognitive development and behavior problems, although these effects may be outweighed by later benefits of having an employed mother [2-4]. There is much less research, however, on the effects of maternal employment during the first year on the physical health of infants and on maternal investments in infant health.

In particular, there is only limited, conflicting information about how maternal employment affects breastfeeding decisions. Although there is some evidence that maternal employment is associated with a reduced duration of breastfeeding among mothers who have initiated breastfeeding, this association may be confounded by unobserved factors that affect both breastfeeding duration and maternal employment decisions, such as cultural beliefs about breastfeeding and the level of family stress. Moreover, little is known about how decisions about employment after childbirth affect breastfeeding initiation. Breastfeeding initiation is an important outcome to study, as a sizeable proportion of mothers (32% in 2000) do not initiate breastfeeding at all [5].

The Surgeon General, in a report on breastfeeding published by the Department of Health and Human Services in 2000, stated that breastfeeding is "…one of the most important contributors to infant health [6]." Breastfeeding is linked to reductions in respiratory illnesses, gastrointestinal illnesses, ear infections, asthma and other allergies, Sudden Infant Death Syndrome, and childhood acute leukemia [7-11]. However, breastfeeding also is a time-intensive activity that requires mothers either to be with their children at every feeding, or to express and store breastmilk for later use. Although many working mothers do feed their children breastmilk, returning to work, particularly full-time work, shortly after the birth of a child may make establishing and continuing with breastfeeding difficult.

This study is based on the hypothesis that among mothers who were employed before the child was born, returning to work before the child is 3 months old (1) reduces the probability of initiating breastfeeding and, (2) among mothers who initiated breastfeeding, reduces the number of weeks of breastfeeding. We also hypothesize that among mothers who return to work within 3 months after childbirth, working full-time rather than part-time is associated with lower rates of initiating breastfeeding and shorter breastfeeding durations. Unlike previous researchers, we are able to take advantage of data on siblings, which allow us to examine breastfeeding and employment decisions within mothers over time. Although this method has its own limitations, we avoid the problem of needing to find good identifying variables, such as valid instruments in an instrumental variables context.

The findings suggest that maternal employment is negatively associated with both breastfeeding initiation and breastfeeding duration. Baseline probit models and fixed effects models indicate that returning to work within 3 months is associated with a reduction in the probability that the mother will initiate breastfeeding by 16-18%. Among those mothers who initiate breastfeeding, returning to work within 3 months is associated with a reduction in the length of breastfeeding by 4-6 weeks. This reduction may have significant health effects since the median number of weeks of breastfeeding among mothers who initiated breastfeeding in our sample was only 12 weeks. We find less consistent evidence that working at least 35 hours per week (among mothers who return to work within 3 months) detracts from breastfeeding. Baseline and fixed effects models indicate that returning to full-time work is associated with a reduction in the length of breastfeeding of 1-4 weeks. However, we do not find consistent evidence regarding the association between returning to full-time work and breastfeeding initiation.

2. BACKGROUND

The American Academy of Pediatrics recommends that mothers breastfeed exclusively (without the use of infant formula or other liquids) for the first six months, and then continue partial breastfeeding for at least a year [12]. In 2000, about 68 percent of mothers in the United States initiated breastfeeding in the hospital, and 31 percent were still breastfeeding (exclusively or partially) when their infants were six months old. These rates are much higher than they were even ten years prior, when only 52 percent of mothers initiated breastfeeding and 18 percent breastfed for at least 6 months [5]. However, breastfeeding rates among African-American mothers and mothers who participate in WIC still lag far behind the Healthy People 2010 objectives of increasing the breastfeeding initiation rate to 75% and the rate of breastfeeding for 5-6 months to 50% [13]. In 2000, 51 percent of African-American mothers initiated breastfeeding in the hospital and 12 percent were still breastfeeding at six months; these rates were 57 percent and 20 percent respectively for WIC participants in 2000 [5].

There is some empirical evidence that maternal employment may interfere with breastfeeding. Visness & Kennedy (1997), for example, use 1988 data from the National Maternal and Infant Health Survey and find that although returning to work within a year of the child's birth is not associated with breastfeeding initiation, returning to work is associated with shorter duration of breastfeeding among those who initiate breastfeeding [14]. Similarly, Lindberg (1996), using Cycle IV of the National Survey of Family Growth, find that part-time working mothers are more likely to initiate breastfeeding and have longer breastfeeding durations compared to full-time working mothers [15]. Fein & Roe (1998) also report that expectations of full-time work during pregnancy and full-time work three months after childbirth are associated with reductions in breastfeeding initiation and duration, respectively. Their work is based on the Infant Feeding Practices Study, conducted in 1993 [16]. As the authors of these papers acknowledge, the observed negative relationship between returning to employment and breastfeeding may be confounded by reverse causality (e.g. the mother returns to work because she has stopped breastfeeding) or by other unobserved factors, such as mothers' unobserved desire to wean the child.

Roe et al. (1999), however, find the same result after using methods that account for the simultaneity of the employment and breastfeeding decisions [17]. These authors estimate a simultaneous model of maternal employment and infant feeding, using maternal occupation to identify the breastfeeding equation. They limit the analysis to mothers who have initiated breastfeeding, who were employed before the child's birth, and who planned to return to work before the child was a year old (although the analysis includes women who changed their plans). The results suggest that the duration of maternal leave is positively associated with the duration of breastfeeding, and working 8 hours per day reduces the number of daily breastfeeding sessions.

These results are based on the identification assumption that maternal occupation affects return to work decisions but does not have direct impact on breastfeeding decisions. Although the authors find that maternal occupation is not a significant predictor of breastfeeding practices, certain occupations, particularly professional occupations where personal space for expressing milk is available and many mothers breastfeed, may encourage a new mother to initiate or continue to breastfeed. Work environment is likely to affect the ability to continue breastfeeding and possibly the decision to initiate breastfeeding. Visness & Kennedy, in fact, report that among white women, occupational category is associated with breastfeeding duration after controlling for the duration of maternity leave. It is difficult to provide a strong argument for any personal or family characteristics being a clearly good predictor of return-to-work decisions, but not associated with breastfeeding decisions. In this study, because we have data on siblings, we avoid the problem of needing to find suitable identifying variables.

3. THEORETICAL MOTIVATION

After the child's birth, mothers maximize the following multi-period, discrete time objective function:

$$\sum_{t=0}^{T} U(H_t, h_t, \alpha_t, \beta_t)$$

subject to the following constraints and identities

$$H_t = H_{t-1} + \mathrm{f}_t(\gamma_t, \delta_t, \varepsilon_t)$$

$$H_0 = H^*$$

$$h_t = h_{t-1} + \mathrm{g}_t(\delta_t, \phi_t)$$

$$h_0 = h^*$$

$$Z = \beta_t + \gamma_t + \delta_t + \eta_t$$

$$\sum_{t=0}^{T} (E_t + \eta_t w) = \sum_{t=0}^{T} (p\alpha_t + q\varepsilon_t + r\phi_t) + K + M$$

The utility in each period is a function of the mother's health (*H*), the child's health (*h*), the mother's consumption (α), and the mother's leisure (β). The mother's health in a period is a function of her health in the last period with adjustments for the time she invests in health producing activities in a given period (γ), the time she invests in her child's health in a given period δ, and the health care expenditures she makes (ε). A mother has an initial health endowment of H^* at $t = 0$.

Three additional points are notable with respect to δ. First, we are characterizing the time investment in the child's health as only including breastfeeding, and breastfeeding can affect the mother's health [18]. Second, we view the time invested in child health (i.e., time spent breastfeeding) as being time above and beyond the time it would take for the mother to use formula to feed the child. Although breastfeeding is not necessarily more time-consuming than formula-feeding, we assume that breastfeeding requires more *maternal* time than formula-feeding (e.g. time spent expressing milk, fewer opportunities for another caregiver to feed child). Third, to simplify the model, we consider only exclusive breastfeeding.

The child's health in a period is a function of the child's health in the last period with adjustments for the time that the mother invests in the child's health in a given period (δ) and the medical care expenditures for the child (ϕ). The initial health endowment for the child is h^* at $t = 0$. Both f and g are subscripted because the investment in the child's health through breastfeeding has a dose response relationship and the marginal impact of breastfeeding on the mother's and child's health will change over time. The dose response relationship suggests an increasing marginal impact of breastfeeding on infant health when the child is very young, but, eventually, the marginal impact will diminish and the marginal benefits will be lower than the marginal costs of continuing to breastfeed.

The fifth constraint indicates that the total time available in a period (Z) is the sum of four choice variables: time spent investing in the mother's health, time spent investing in the child's health, time spent in leisure, and time spent working (η). The final constraint is a lifetime budget constraint with an interest rate of 0—having any other interest rate would complicate the expression without adding significantly to the insight. This is not meant to disallow saving and borrowing—only to simplify the expression. The left hand side represents resources available: the sum of endowments (or non-labor income) plus the wage multiplied by the time spent working in each period. The right hand side represents expenditures. The first three terms that are included in the summation over time are straightforward. The price of consumption goods is p, the price of medical care for the mother is q, and the price of medical care for the child is r. K and M require additional explanation

The decision to breastfeed at all and the decision on how long to breastfeed include discrete costs that are only incurred if the mother decides to breastfeed (K=0 if $\delta_t = 0 \,\forall\, t$) and only if the mother decides to breastfeed while working (M=0 if ($\delta_t = 0$ or $\eta_t = 0$) $\forall\, t$). The decision to initiate breastfeeding involves start-up costs (K) above and beyond start-up costs associated with formula feeding, such as purchasing nursing clothing and spending time and effort learning about breastfeeding. The decision to breastfeed while working requires additional start-up costs (M) such as the purchase of a breast pump and finding appropriate time and space to express milk. The "startup" costs of breastfeeding are described only as financial in this case, but they could be both financial and psychological.

Without these two discrete costs, the mother would have six choice variables since η_t (work time) is implied by the other choices and the expression can be rewritten without η_t. The choices include the time to invest in the mother's health (γ), the time to invest in the child's health (δ), the mother's medical care (ε), the child's medical care (ϕ), leisure (β), and consumption (α). As with any economic optimization problem, the mother will choose to breastfeed until the marginal costs (in terms of the opportunity cost of time when the expression is rewritten without η_t) are just equal to the marginal benefits (in terms of the effects on the mother's health, the effect on the child's health, and savings on formula expenditures).

The two discrete costs allow corner solutions. First, a mother might choose to stop breastfeeding when she returns to work rather than continuing to breastfeed longer if the fixed startup cost of trying to breastfeed and work (M) is higher than the net marginal utility from continuing to breastfeed after returning to work. In that case, the mother evaluates whether or not to initiate breastfeeding by comparing two constrained optimization problems: breastfeeding until she returns to work ($\delta_t = 0 \,\forall\, t$ where $\eta_t > 0$) with a cost of K for initiating breastfeeding and no breastfeeding ($\delta_t = 0 \,\forall\, t$).

Thus, while returning to work is a decision that is a choice variable similar to the choices to initiate and continue to breastfeed, the discrete cost of trying to work and breastfeed leads to a potentially clear directional relationship between the decision to return to work and to breastfeed. A decision to return to work may limit the time spent breastfeeding. Furthermore, if the incremental net benefit of breastfeeding (excluding *M*) that would occur after the return to work were necessary to offset some part of *K* to make it worthwhile to breastfeed at all, then the decision to return to work could also affect the decision on whether or not to breastfeed at all.

4. METHODS

This paper is based on the hypothesis that among mothers who were employed before the birth of a child, returning to work, particularly full-time work, shortly after birth reduces the initiation and duration of breastfeeding. To test this hypothesis, the study focuses on estimating the following equation:

$$1)\ BF_{ij} = b_0 + b_1E_{ij} + b_2X_i + b_3Y_j + u_i + e_{ij}.$$

This equation is specific to the ith child of mother j. Two dependent variables (BF) are used in this analysis: (1) a dichotomous indicator of whether or not the mother initiated breastfeeding; and (2) among mothers who initiated breastfeeding, the log of the number of weeks the child was breastfed. We have complete information on all breastfeeding spells, which eliminates the problem of censored information about the dependent variable.

The main independent variable of interest is maternal employment status after the birth of a child (E). We are primarily interested in two aspects of employment status: (1) the age of the child in weeks when the mother returned to work; and (2) among mothers who had returned to work by the time the child was three months old, the number of hours worked per week. We consider both the timing of return to work and the intensity of the work effort because previous research suggests that both of these factors may interfere with breastfeeding. To measure employment status after birth (E_{ij}), we use a dummy variable indicating whether or not the mother returned to work before the child was three months old. For mothers who returned to work within three months, we consider a dummy variable indicating whether or not the mother worked full-time (at least 35 hours per week).[2, 3]

The vector X_i includes observed child-specific factors that may determine breastfeeding, such as the child's gender and health endowment at birth, as proxied by low birth-weight, as well as the child's birth order and year of birth. The child's year of birth is included to capture secular trends in infant feeding practices. The vector Y_j includes observed mother-specific factors that may determine breastfeeding initiation and duration. These factors were selected based on previous literature and include education, age, family size, Armed Forces Qualification Test (aptitude test) score, marital status, family income, region of residence and whether or not the family participated in the Food Stamp program. These variables are measured during the year of the child's birth.

The models also include a measure of maternal smoking during pregnancy. It is not expected that maternal smoking will directly affect the decision to breastfeed. However, smoking during pregnancy may proxy the mother's unobserved motivation to make investments in infant health, such as breastfeeding. Maternal smoking, along with low birth weight and some other right hand side variables, may be endogenous in the model because they may be correlated with unobserved factors that are associated with both breastfeeding and employment decisions. For this reason, all of the models also were estimated with a parsimonious set of covariates (child's race, year of birth, and mother's aptitude test score).

Even with a rich secondary data set, there may exist important, unobserved factors that affect breastfeeding decisions and that are not adequately captured by the available data. This

[2] Our results are not sensitive to changing the full-time work threshold to 40 hours per week.
[3] We also estimate these models using a continuous measure of the child's age in weeks when his/her mother returned to work and a continuous measure of the number of hours the mother worked per week when she returned to work. In the paper, we do not present results from these analyses. These results are qualitatively very similar to those presented here and are available upon request.

possibility is particularly relevant to this paper because the data set used in the study is a labor market survey, which was not intended to focus on the determinants of infant feeding decisions. The vector u_i represents time-invariant, unobserved maternal factors that affect breastfeeding. These factors may include the mother's cultural beliefs about breastfeeding, her mental and physical health, her knowledge and interest in child health, and her social support system.

It is possible that some of these unobserved maternal factors that affect breastfeeding also are correlated with unobserved factors that determine maternal employment after the birth of a child. It is also possible that the employment and breastfeeding decisions are made simultaneously, with breastfeeding decisions affecting employment decisions and employment decisions affecting breastfeeding decisions. For example, mothers who do not intend to breastfeed their infants might be more likely than mothers who intend to breastfeed to return to work shortly after the birth. Because intention to breastfeed is unobserved, standard estimation of equation 1 would lead to a biased estimate of the impact of maternal employment on breastfeeding, since, in this scenario, mothers who return to work shortly after the birth are more likely than other mothers to not intend to breastfeed.

Initially, ordinary least squares (OLS), and standard probit models are used to estimate equation 1. Estimating equation 1 by OLS or a standard probit, however, can lead to biased and inconsistent estimates if these endogeneity problems exist. This study uses two approaches to address this problem. First, observed data on maternal characteristics is used to proxy u_j to the fullest extent possible, and we use informal and formal tests to assess whether or not the problem of unobserved heterogeneity is important. The informal test consists of comparing results from parsimonious and more fully specified models of breastfeeding. If controlling for observed heterogeneity affects the estimated impact of maternal employment on breastfeeding, it seems likely (although we can never know for certain) that unobserved heterogeneity also is important in this relationship. More formally, we also use the Breusch-

Pagan Lagrange Multiplier test to determine whether or not the disturbance term includes a group-level component.

These tests suggest that unobserved heterogeneity may be important in the relationship between maternal employment and breastfeeding. Therefore, we attempt to account for unobserved mother-specific factors by estimating family-level fixed effects models. These models take advantage of the fact that most mothers in the data have multiple children. Consequently, we have repeated observations on employment and breastfeeding decisions for much of the sample. Ideally, we would have used random effects models to account for a random, unobserved mother-specific factor that affects breastfeeding and employment decisions. However, a Hausman test rejected the null hypothesis that the right hand side variables in the model were exogenous with respect to the error term, which indicates that the random effects model would not lead to consistent estimates of the parameters.

Fixed effects models are robust to correlation between right-hand side variables and the disturbance term. These models also address the problem of time-invariant, mother-specific factors, such as the mother's fixed, cultural beliefs about breastfeeding. However, because these models are based on "within" variation only, we cannot estimate the impact of time-invariant factors, such as race and aptitude test score, on breastfeeding. Moreover, we cannot account for unobserved, mother-specific factors that might change between the birth of two children, such as her level of family stress. It is possible that the causation runs the opposite way, with mothers returning to work later because they are breastfeeding. Roe et al. find little empirical support for this direction of causality, but it remains a possibility. The fixed effects models will not eliminate this simultaneity problem if the unobserved factor that governs both decisions, such as the intention to breastfeed, is time-varying. This issue suggests that these findings should be interpreted and generalized with caution.

A Breusch-Pagan test based on a standard fixed effects model shows evidence of heteroscedasticity in the random component of the error term. For this reason, a heteroscedastic fixed effects model is estimated using feasible generalized least squares. This model allows for family-specific variances, but not for cross-sectional correlation within panels. A standard fixed effects model led to very similar results.

This study does not use the instrumental variables (IV) method or the bivariate probit model to address the problem of unobserved heterogeneity. These methods do have certain advantages over the fixed effects models, but practical implementation of these methods requires that we have at least one exogenous variable that is a predictor of the maternal employment decision after the birth but is not correlated with either breastfeeding or the error term in the breastfeeding equation. Unfortunately, there are few, if any, characteristics of the mother or family that would be appropriate as identifying instruments. Roe et al. (1999), in their study of maternal employment and breastfeeding, identify their breastfeeding equation using occupational category as an identifying instrument. It seems likely, however, that occupational category also affects breastfeeding decisions since some occupations (e.g. professional) offer more flexibility and personal space than other occupations (e.g. clerical). Although no national survey of workplace attitudes towards breastfeeding exists, small-scale surveys indicate that employers generally do not view providing breastfeeding support as a priority and attitudes towards breastfeeding vary by employer characteristics such as firm size [19-20].

State-level policies are more likely than individual-level characteristics to be exogenous. However, state-level policies are likely to be very poor predictors of individual mothers' decisions to return to work. Bound et al. (1995), Bollen et al. (1995), Nelson & Startz (1990), Staiger & Stock (1994) and others all have noted that a low first stage F-statistic for the identifying instrumental variables may suggest that the TSLS estimates are no better than biased OLS estimates [21-24].

5. DATA

Data used in the study come from the 1998 releases of the National Longitudinal Survey of Youth (NLSY79) and the Children of the National Longitudinal Survey of Youth (CoNLSY). NLSY79 is an annual, national survey that was initiated in 1979 with a sample of 12,686 young people who at that time were aged 14-21. The original sample includes a nationally representative sample of civilian youth as well as over-samples of African-Americans, Hispanics, indigent Whites, and Armed Forces personnel. Children of NLSY79 mothers (CoNLSY respondents) are interviewed and/or assessed in a separate, linked survey. We limit the sample to children born before 1997 in order to have data on completed breastfeeding spells for all children. As of 1996, 7,103 CoNLSY respondents ranging in age from infant to over 21 years old were assessed and/or interviewed. Of these 7,103 respondents, 21% were Hispanic, 33% were African-American, and 46% were White/Other.

Two analysis samples are used in this study: (1) the full sample (N=5,804), which includes children born between 1974 and 1996 to mothers *who were employed full-time or part-time at some point during the year that preceded the child's birth*; and (2) the sibling sample (N=3,947) which limits the full sample to children who have at least one sibling in the sample. The sibling sample is used to estimate family fixed effects models. We limit both samples to children born to mothers who were employed before the birth of the child in order to focus on how the timing and intensity of return to work affects breastfeeding decisions. When we focus on the intensity of return to work, we limit the sample to children of mothers who returned to work within three months after the birth of the child. When we analyze breastfeeding duration as an outcome, we limit the sample to children whose mothers initiated breastfeeding.

Both samples exclude observations with missing data on breastfeeding or missing data on the mother's employment status during the year before or during the year after the child's birth. The samples include respondents with missing data on Food Stamps participation, birth weight, education, family income, marital status, family size, AFQT score, and smoking during pregnancy. For these respondents, missing variables are replaced by unconditional means. In order to check the sensitivity of the results to this imputation, the models were re-estimated after dropping respondents with missing information. The results were almost identical to those presented here, and suggest that the results are not sensitive to the imputation method used to deal with missing data.

6. DESCRIPTIVE STATISTICS

Table 1 presents means and standard deviations for the full and sibling samples. In both samples, about 23% of the respondents are African-American and about 19% are Hispanic. The average maternal age is 25, and the average number of years of education is 12.6. Most mothers (about 70%) were married at the time of the birth, and about 31% reported smoking during pregnancy. Employment after the birth of the child was common in both samples, with 55-59% of mothers returning to employment within 3 months after the child's birth, but only 34-36% of the samples returned to full-time work before the child was 3 months old. Klerman & Leibowitz using 1986-1988 data from the Current Population Survey find that about 40 percent of new mothers return to work within three months, but their sample includes part-time workers and mothers who were not working before the child's birth [25]. (The mean year of birth in our sample is 1986.) About 51% of each sample initiated breastfeeding, and, among those who initiated, the mean duration of breastfeeding was 18-19 weeks.

Table 2 presents means and standard deviations by the mother's employment status after the child's birth. Columns 1-2 compare maternal characteristics between mothers who returned to work before the child was 3 months old and mothers who did not. Columns 3-4 compare maternal characteristics between mothers who returned to work full-time within 3 months after the birth and mother who had returned to work within 3 months, but not full-time.

The timing of return to work (Columns 1-2) appears to be associated with breastfeeding duration, but not initiation. Mothers who returned to work within 3 months were just as likely to initiate breastfeeding, but they breastfed for about 3.5 fewer weeks than mothers who did not return to work within 3 months. There were other interesting differences between the two groups. Mothers who returned to work within 3 months were older, more educated, more likely to be married, had higher family income and were less likely to have smoked during pregnancy compared to mothers who did not return to work within 3 months. These differences were consistent, but less striking, when the comparison is made between full-time working mothers and mothers who were working within 3 months, but were not working full-time (Columns 3-4). Working full-time, however, appears to be associated with decreased likelihood of breastfeeding initiation and a decline of about 5 weeks of breastfeeding. These findings suggest that the small differences in breastfeeding between these groups of mothers may be confounded by the higher socioeconomic status of mothers who return to work shortly after childbirth.

Table 3 displays a more detailed breakdown of breastfeeding patterns by return-to-work status. As the top panel of Table 3 shows, mothers who return to work within a month after childbirth are only slightly less likely to initiate breastfeeding than mothers who return later. Mothers who return to work when the child was 1-3 months old, however, have lower rates of breastfeeding initiation than mothers who did not return to work within 3 months. There was no consistent pattern between the timing of return to work and breastfeeding duration. As the bottom panel of Table 3 demonstrates, there is a much stronger relationship between hours worked and breastfeeding among mothers who had returned to work within 3 months after childbirth. The number of hours worked is inversely associated with both breastfeeding initiation and duration. These results suggest that the timing of return to work may or may not be important in breastfeeding decisions, but among mothers who return to work within 3 months, working more hours detracts from breastfeeding. Clearly, other observable and unobservable factors may confound these relationships.

Fixed effects models rely solely on within-family variation in employment and breastfeeding practices since all other variation is eliminated from the model. Tables 4a and 4b assess whether or not there is sufficient within-family variation in the data. For clear presentation, the tables focus only on breastfeeding initiation and only on families with 2 children. We note, however, that there is much more scope for variation within larger families (which essentially have more repeated observations) and when the number of weeks of breastfeeding is considered as an outcome. Our assessment of within-family variation therefore focuses on the scenario where within-family variation is likely to be the lowest. Approximately 67% of the children in the sibling sample lived in 2-child families.

Table 4a shows variation within 2 child families in returning to work within 3 months, initiating breastfeeding, and returning to full-time work within 3 months. About 35% of mothers returned to work within 3 months with one child, but did not return within 3 months with another child. Approximately 31% of mothers worked full-time before one child was 3 months old, but did not work full-time within 3 months with another child. Variation in breastfeeding initiation within families was less common; only about 19% of mothers initiated breastfeeding with one child but not with another child.

Table 4b limits the comparison to mothers who show variation between two children in both breastfeeding initiation and work decisions after childbirth. These "switchers" actually determine the estimate of maternal employment on breastfeeding in the fixed effects models. Within 90 switcher families, about 41% of mothers initiated breastfeeding when the mother returned to work within 3 months, while about 59% of mothers who did not return to work within 3 months initiated breastfeeding. However, within switcher families, full-time work did not appear to negatively affect breastfeeding decisions. The fixed effects models presented in the next section enhance this simple comparison by controlling for time-varying and time invariant covariates, such as the mother's marital status and the birthweight of the child, that may confound these relationships.

Regression diagnostics revealed that one of the dependent variables, the number of weeks the child was breastfed, was highly right skewed. Once this variable was transformed to its natural log form, several regression problems were solved at once, including a problem of non-normal error distribution and several influential data points. For this reason, all weeks breastfed models were estimated in semi-log form.

7. ESTIMATION RESULTS

Tables 5 and 6 show estimates of the timing of return to work (Table 5) and the intensity of return to work (Table 6) on breastfeeding initiation and breastfeeding persistence. Columns 1-3 in each table focus on the initiation of breastfeeding as an outcome, while columns 4-6 in each table show results from models in which the log of the number of weeks breastfed is the dependent variable. The weeks breastfed models are limited to children whose mothers initiated breastfeeding. All of the models in Table 6, which focuses on the intensity of work, are limited to children whose mothers returned to employment before they were three months old.

Timing of Return to Work and Breastfeeding

Column 1 in Table 5 shows results from a parsimonious model of breastfeeding initiation. This model includes as independent variables a dummy indicator of whether or not the mother returned to work before the child was 3 months old, race, the year of the child's birth, and the mother's aptitude test score. Column 2 expands this model to include a larger number of child and maternal characteristics, some of which are potentially endogenous. In both the parsimonious and fully specified models, returning to work within 3 months is associated with a decrease in the probability of .08 -.09 in the probability of initiating breastfeeding. At the sample means, this reduction translates into a 16-18 percent decrease in the probability of initiating breastfeeding.

The inclusion of additional covariates (Column 2) has little impact on the magnitude of the estimated impact of returning to work on breastfeeding initiation. Many of the additional covariates, however, appear to be important predictors of breastfeeding initiation. Older, more educated, and married mothers are more likely than others to initiate breastfeeding. African-American and Hispanic mothers are less likely than mothers of other races and ethnicities to initiate breastfeeding. Children who are first-born are more likely to be breastfed, and low birthweight children are much less likely to be breastfed. Mothers who report smoking during pregnancy are less likely to initiate breastfeeding. It appears that selection along unobserved characteristics may not be particularly important because: (1) observed maternal and child characteristics are good predictors of breastfeeding initiation; and (2) the inclusion of these characteristics does not have much impact on the estimate of maternal employment on breastfeeding initiation.

If selection on observed characteristics is not particularly important, it seems unlikely that selection on unobserved characteristics is important [26]. This informal test suggests that controlling for unobserved heterogeneity may not be important in this context, but it is by no means a definite indication of the importance of unobservable factors. In this case, in fact, a Breusch-Pagan Lagrange Mulitplier test for random effects suggested that the disturbance term includes a mother-specific error component. This finding is consistent with breastfeeding research, which suggests that mother-specific factors that are difficult to measure in a secondary data set, such as a mother’s cultural beliefs about breastfeeding, are very important in the decision to initiate breastfeeding. For this reason, we estimate a fixed effects model using a sub-sample of the data that only includes children of mothers with more than one child. This model differences out the fixed, mother-specific component of the error term, leaving an unbiased estimate of the impact of returning to work on breastfeeding initiation.

Column 3 of Table 5 shows fixed effects model results. The models ignore the fact that the dependent variable is binary. A random effects probit would have accounted for the binary dependent variable, but a Hausman test suggested the existence of correlation between right hand side variables and the disturbance term. For this reason, a standard fixed effects model was estimated. (It was not possible to test for heteroscedasticity because of the large number of mothers in the sample.) The fixed effects model results are consistent with the baseline probit results – returning to work within 3 months is associated with a reduction in the probability of breastfeeding. It is difficult to interpret the magnitude of this estimate, however, since the binary nature of the dependent variable was ignored.

Columns 4, 5 and 6 present results of models in which the log of the number of weeks breastfed is the dependent variable. These models are limited to children whose mothers initiated breastfeeding. Columns 4 and 5 present findings from a parsimonious OLS model and a fully specified OLS model. As before, the inclusion of a full set of covariates does not change the magnitude of the coefficient on maternal employment very much. In both models, returning to work before the child is 3 months old is associated with a 33-34% decrease in the number of weeks of breastfeeding, among children whose mothers initiated breastfeeding. At the sample means, this percentage decrease translates into 6 fewer weeks of breastfeeding.

Older, more educated, married mothers breastfed their infants longer than younger, less educated, unmarried mothers. First-born children actually were breastfed for fewer weeks than later-born children, even though previous results showed that mothers were more likely to initiate breastfeeding with first-born children. Mothers who smoked during pregnancy breastfed their infants for fewer weeks compared to mothers who did not smoke during pregnancy. As in the breastfeeding initiation models, maternal smoking during pregnancy may represent a mother's lack of knowledge or interest in child health.

As before, a Breusch-Pagan Lagrange Multiplier test supported the existence of a fixed, mother-specific component in the disturbance term, and a Hausman test suggested that fixed effects models rather than random effects models are appropriate in this case. In the breastfeeding duration models, however, the sample size was reduced because the models were limited to children whose mother initiated breastfeeding. The smaller sample size allowed us to perform a Breusch-Pagan test for heteroscedasticity in the standard fixed effects model. Because we did find evidence of heteroscedasticity, we estimated a heteroscedastic fixed effects model, which allows for mother-specific variances.

Findings from the heteroscedastic fixed effects model support the baseline results. Returning to work within 3 months is associated with a 22% decrease in the number of weeks of breastfeeding, among mothers who initiated breastfeeding. At sample means, this percentage decrease represents a reduction in breastfeeding of about 4 weeks. In summary, then, all of the models support the idea that returning to work within 3 months of childbirth is associated with a reduced probability of initiating breastfeeding and, among mothers who initiate breastfeeding, a reduction of 4-6 weeks in the duration of breastfeeding.

Intensity of Return to Work and Breastfeeding

It is assumed that returning to work interferes with breastfeeding because both activities are time-intensive. Because working full-time is clearly more time-intensive than working part-time, it is possible that among mothers who return to work within 3 months, working more hours may have more detrimental effects on breastfeeding than working fewer hours. The descriptive statistics presented in Table 3 show that full-time, working mothers are less likely to breastfeed than part-time working mothers. However, in Table 4b, there does not appear to be a clear trend in the rate of breastfeeding within mothers who had different working intensity between 2 children.

Table 6 presents results from models that are limited to mothers who returned to work within 3 months after childbirth. Columns 1-3 show findings from models of breastfeeding initiation, while columns 4-6 show results from models with log of the number of weeks of breastfeeding as the dependent variable. All of the baseline results (columns 1-2 and 4-5) suggest that among mothers who returned to work within 3 months, working full-time reduces the probability of breastfeeding by about 18% and reduces the number of weeks of breastfeeding by about 4 weeks (among those who initiated breastfeeding). The comparison group includes all mothers who returned to work within 3 months, but were working less than 40 hours a week. The median hours worked in the comparison group was 25 hours per week.

The fixed effects models (Table 6, columns 3 and 6) do not completely support these results. The models suggest that working full-time may have a negative impact on the number of weeks of breastfeeding of about 1 week among mothers who initiate breastfeeding (column 6). This result is consistent with our previous findings. However, after differencing out an unobserved, mother-specific effect, working full-time appears to have a positive association (rather than a negative association) with breastfeeding initiation (column 3). This positive result is inconsistent with both the theoretical motivation and all of the other empirical results in this paper.

8. DISCUSSION AND CONCLUSIONS

Breastfeeding is one of the most important maternal investments in infant health during the first year of life, but very few health economics studies focused on the United States consider breastfeeding as a health behavior that can be influenced by labor market decisions and by public policy. This study provides evidence that return to work decisions after childbirth can have important effects on breastfeeding. Therefore, policies that affect return to work decisions after childbirth, through their effects on breastfeeding, have considerable potential to improve infant health. We find that returning to work within the first three months of the infant's life is associated with a reduction in the probability of initiating breastfeeding by 16-18 percent, and a reduction in the duration of breastfeeding by 4-6 weeks among mother who initiate breastfeeding. Full-time work among mothers who return to work within 3 months has no consistent effect on breastfeeding initiation, but it is associated with a reduction in the duration of breastfeeding of about 1-4 weeks.

Our results are novel and policy-relevant for several reasons. First, we use data on siblings, which allows us to build on previous literature on employment and breastfeeding by addressing the problem of unobserved heterogeneity without relying on potentially weak identifying variables. Past studies, with the exception of Roe et al., ignore the problem of unobserved heterogeneity. Our results on the effects of return to work decisions on breastfeeding duration are consistent with previous work, which suggests that the length of maternity leave and the number of hours worked are positively associated with breastfeeding duration.

We also consider the effect of return-to-work decisions on breastfeeding initiation. This outcome is important to address because: (1) a sizable proportion of mothers do not initiate breastfeeding; (2) as we describe, theory suggests that returning to work can affect initiation; and (3) most previous research focuses on the effect of employment on breastfeeding duration among mothers who have already initiated breastfeeding. We find that returning to work within 3 months detracts from breastfeeding initiation. Because the decision to breastfeed is made immediately after childbirth, we interpret this association as a mother's plans to return to work within 3 months (which we cannot measure directly) having a negative association with breastfeeding initiation. Plans to return to full-time work rather than part-time work do not appear to have a consistent association with breastfeeding initiation.

Finally, the magnitude of the associations we find are large and important from a public health perspective. A reduction in breastfeeding duration of six weeks implies that, at the mean, a mother would stop breastfeeding her child when he is between two and three months old instead of when he is about four months old. Clinical literature suggests a dose response relationship between the length of breastfeeding and health benefits for infants such as decreased otitis media, acute respiratory infection, diarrhea, gastroenteritits, and allergic disease [27-30], although these findings are not universal [31]. Our results suggest, therefore, that decisions about employment after childbirth, through their impact on breastfeeding, may have important ramifications for infant health.

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Table 1: Sample Means and Standard Deviations		
	Full Sample (N = 5,804)	Sibling Sample (N = 3,947)
Initiated breastfeeding	0.505 (0.500)	0.512 (0.500)
Number of weeks breastfed (if initiated)	18.373 (19.864) N=2,749	19.07 (20.05) N=1,897
Child is first born	0.486 (0.500)	0.375 (0.482)
Child is female	0.496 (0.500)	0.494 (0.500)
Year of child's birth, 19--	86.17 (4.55)	86.28 (4.44)
Revised AFQT score percentile	40.08 (26.36)	41.15 (26.40)
African-American	0.235 (0.424)	0.230 (0.421)
Hispanic	0.186 (0.389)	0.196 (0.397)
Age of mother	25.25 (4.52)	25.37 (4.47)
Lived in South	0.367 (0.482)	0.353 (0.478)
Lived in Central	0.239 (0.427)	0.252 (0.434)
Lived in West	0.200 (0.400)	0.210 (0.408)
Received Food Stamps	0.177 (0.358)	0.173 (0.356)
Low birthweight	0.080 (0.266)	0.078 (0.262)
Family size	3.49 (1.60)	3.63 (1.62)
Mother's education in years	12.60 (2.11)	12.64 (2.10)
Mother is married	0.698 (0.442)	0.718 (0.434)
Net family income in dollars	33,114 (62,974)	33,605 (64,303)
Smoked during pregnancy	0.319 (0.453)	0.306 (0.446)
Returned to work before child was 3 months old	0.550 (0.498)	0.586 (0.492)
Worked at least 35 hours a week before child was 3 months old	0.345 (0.475)	0.357 (0.479)

Table 2: Sample Means and Standard Deviations by Maternal Employment Status Main Sample (N = 5,804)				
	(1) Returned to work before child was 3 months old (N=3,192)	(2) Did not return to work before child was 3 months old (N=2,612)	(3) Worked at least 35 hours a week before child was 3 months old (N=2,369)	(4) Did not work at least 35 hours a week before child was 3 months old (N=3,435)
Initiated breastfeeding	0.506 (0.500)	0.505 (0.500)	0.483 (0.500)	0.521 (0.500)
Number of weeks breastfed among those who initiated	16.52 (18.47) (N=1531)	20.71 (21.27) (N=1218)	14.91 (16.56) (N=1,089)	20.64 (21.47) (N=1,660)
Child is first born	0.464 (0.499)	0.511 (0.500)	0.487 (0.500)	0.485 (0.500)
Child is female	0.497 (0.500)	0.495 (0.500)	0.504 (0.500)	0.491 (0.500)
Year of child's birth, 19--	87.16 (4.52)	84.98 (4.30)	87.28 (4.50)	85.42 (4.43)
Revised AFQT score percentile	43.65 (26.47)	35.71 (25.56)	42.60 (26.50)	37.96 (26.14)
African-American	0.229 (0.420)	0.242 (0.429)	0.260 (0.439)	0.218 (0.413)
Hispanic	0.182 (0.386)	0.191 (0.393)	0.192 (0.394)	0.183 (0.386)
Age of mother	26.26 (4.38)	24.01 (4.37)	26.41 (4.29)	24.45 (4.50)
Lived in South	0.384 (0.486)	0.347 (0.476)	0.417 (0.493)	0.333 (0.471)
Lived in Central	0.242 (0.429)	0.235 (0.424)	0.215 (0.411)	0.256 (0.437)
Lived in West	0.179 (0.383)	0.226 (0.418)	0.180 (0.384)	0.214 (0.410)
Received Food Stamps	0.115 (0.285)	0.254 (0.419)	0.107 (0.274)	0.226 (0.400)
Low birthweight	0.076 (0.258)	0.084 (0.275)	0.084 (0.270)	0.078 (0.263)
Family size	3.47 (1.54)	3.52 (1.68)	3.43 (1.52)	3.58 (1.66)
Mother's education in years	13.00 (2.04)	12.10 (2.10)	13.00 (2.07)	12.27 (2.10)
Mother is married	0.746 (0.414)	0.640 (0.466)	0.741 (0.415)	0.664 (0.456)
Net family income in dollars	36,163 (58,861)	29,388 (67,486)	35,615 (56,417)	30,004 (67,062)
Smoked during pregnancy	0.281 (0.434)	0.366 (0.470)	0.288 (0.437)	0.343 (0.462)

Table 3: Timing of Return to Work, Hours Worked, and Breastfeeding [N]				
	All Mothers Who Were Employed in Year Before Child's Birth			
	Did not return to work within three months after birth	Returned to work when child was less than one month old	Returned to work when child was 1-2 months old	Returned to work when child was 2-3 months old
Initiated breastfeeding	53.6% [1,832]	44.9% [354]	47.8% [609]	45.8% [397]
Mean number of weeks of breastfeeding	18.31 [896]	17.81 [144]	13.34 [276]	15.14 [174]
Mothers Employed Within Three Months After Birth				
	Less than 10 hours per week (N = 152)	Between 10 and 20 hours per week (N=264)	Between 20 and 35 hours per week (N=612)	At least 35 hours per week (N=2,369)
Initiated Breastfeeding	73.0% [152]	58.3% [264]	56.4% [612]	48.3% [2,369]
Number of weeks of breastfeeding	23.1 [141]	23.0 [141]	18.9 [318]	15.2 [1,065]

Tables 4a and 4b: Analysis of Switchers

Table 4a: Variation in Return to Work and Breastfeeding Within 2-Child Families

Families with 2 children, N = 1318			
	Returned to work within 3 mos.	Initiated breastfeeding	Returned to at least 35 hours of work per week within 3 mos.
# of families			
neither child	335 (25.4%)	540 (41.0%)	561 (42.6%)
one child	464 (35.2%)	246 (18.7%)	402 (30.5%)
both children	519 (39.4%)	532 (40.4%)	355 (26.9%)

# of families: number of families with column characteristics, e.g. there are 335 families (25.4% of all two child families) in which neither child was less than 3 months old when the mother returned to work

Table 4b: Two child families with variation in return to work AND variation in breastfeeding between children

	N = 180 children (90 families)		N = 166 children (83 families)	
	Returned to work within 3 months	Did not return to work within 3 months	Returned to at least 35 hours work per week within 3 months	Did not return to at least 35 hours work per week within 3 months
Initiated breastfeeding	37	53	45	38
Did not initiate breastfeeding	53	37	38	45

Table 5
Timing of Return to Work and Breastfeeding

	Initiated Breastfeeding			Log of Number of Weeks Child was Breastfed		
	(1) Probit (marginal effects)	(2) Probit (marginal effects)	(3) Fixed Effects Model	(4) OLS	(5) OLS	(6) Heteroscedastic Fixed Effects Model
Mother returned to work before child was three months old	-0.081 (-5.30)	-0.093 (-5.92)	-0.054 (-3.50)	-0.331 (-7.07)	-0.333 (-7.06)	-0.211 (-40.68)
Female child	0.008 (0.560)	0.013 (0.950)	0.013 (1.05)	0.022 (0.520)	0.029 (0.690)	0.073 (14.00)
Year of child's birth	0.014 (8.60)	0.007 (1.85)	0.031 (1.92)	0.016 (3.09)	-0.016 (-1.35)	0.098 (15.53)
African-American	-0.186 (-8.28)	-0.160 (-6.38)		-0.082 (-1.02)	-0.098 (-1.13)	
Hispanic	0.068 (2.85)	0.005 (0.200)		-0.090 (-1.27)	-0.185 (-2.54)	
Mother's AFQT score	0.006 (14.92)	0.004 (9.09)		0.009 (8.66)	0.007 (5.42)	
Child is first born		0.074 (4.70)	0.082 (5.11)		-0.109 (-2.22)	0.267 (39.84)
Age of mother		0.009 (2.02)	-0.016 (-1.01)		0.029 (2.34)	-0.047 (-7.30)
Low birth-weight		-0.125 (-4.26)	-0.096 (-3.52)		-0.146 (-1.39)	-0.098 (-6.53)
Family size		-0.012 (-2.22)	-0.017 (-3.39)		0.017 (0.940)	-0.015 (-4.78)
Mother's education in years		0.023 (4.15)	0.017 (1.85)		0.030 (2.04)	0.041 (13.67)
Mother is married		0.086 (4.10)	0.024 (1.10)		0.125 (1.75)	0.222 (21.29)
Log Family Income		-0.017 (-1.90)	-0.010 (-1.08)		-0.026 (-0.990)	0.020 (10.15)
Mother Smoked During Pregnancy		-0.064 (-3.34)	-0.032 (-1.30)		-0.170 (-2.88)	0.059 (5.67)
Central region		0.016 (0.590)	-0.034 (-0.590)		0.095 (1.30)	0.144 (6.42)
South region		-0.024 (-1.07)	-0.027 (-0.560)		0.001 (0.010)	0.017 (0.650)
West region		0.198 (7.17)	0.008 (0.150)		0.205 (2.85)	-0.068 (-3.02)
Food Stamps		-0.025 (-1.07)	0.030 (1.30)		-0.146 (-0.180)	0.062 (3.79)
Pseudo R-squared	0.111	0.147	---	0.07	0.09	---
Breusch-Pagan LM test for random effects (test stat and p-value)			835.32 (0.000)			179.00 (0.000)

Hausman specification test (test stat and p-value)			33.50 (0.004)			26.25 (0.036)
Breusch-Pagan test for heteroscedasticity (test stat and p-value)			---			1211.89 (0.000)
N families			1728			674
N children	5804	5804	3947	2692	2692	1507

Note: Huber t-statistics (standard models) and t-statistics (fixed effects models) in parentheses, and intercept not shown.

Table 6
Intensity of Return to Work and Breastfeeding Among Mothers who Returned to Work Within 3 Months After Birth

	Initiated Breastfeeding			Log of Number of Weeks Child was Breastfed		
	(1) Probit	(2) Probit	(3) Fixed Effects Model	(4) OLS	(5) OLS	(6) Heteroscedastic Fixed Effects Model
Worked at least 35 hours a week before child was 3 months old	-0.090	-0.094	0.061	-0.202	-0.231	-0.047
	(-4.56)	(-4.62)	(2.29)	(-3.71)	(-4.01))	(-3.34)
Female child	-0.008	-0.005	0.003	0.055	0.054	0.111
	(-0.470)	(-0.270)	(0.140)	(1.03)	(1.00)	(12.82)
Year of child's birth	0.013	0.010	0.038	0.022	-0.008	0.167
	(6.48)	(2.39)	(1.61)	(3.64)	(-0.640)	(10.49)
African-American	-0.155	-0.129		0.015	-0.029	
	(-6.30)	(-4.56)		(0.170)	(-0.310)	
Hispanic	0.093	0.033		-0.145	-0.237	
	(3.71)	(1.19)		(-1.96)	(-3.05)	
Mother's AFQT score	0.005	0.004		0.009	0.007	
	(13.82)	(8.19)		(8.21)	(4.97)	
Child is first born		0.115	0.082		-0.106	0.306
		(5.34)	(3.56)		(-1.66)	(27.77)
Age of mother		0.005	-0.027		0.028	-0.116
		(1.07)	(-1.13)		(2.06)	(-7.39)
Low birth-weight		-0.113	-0.074		-0.124	-0.109
		(-3.16)	(-1.82)		(-1.08)	(-8.52)
Family size		-0.007	-0.019		0.017	0.018
		(-0.930)	(-2.35)		(0.750)	(2.74)
Mother's education in years		0.025	0.008		0.050	0.014
		(4.18)	(0.610)		(3.08)	(2.06)
Mother is married		0.085	-0.046		0.086	0.120
		(3.26)	(-1.31)		(1.02)	(3.18)
Log Family Income		-0.015	0.002		-0.046	0.049
		(-1.27)	(0.110)		(-1.32)	(7.52)
Mother Smoked During Pregnancy		-0.071	-0.055		-0.161	0.067
		(-3.20)	(-1.46)		(-2.32)	(5.06)
Central region		0.016	0.045		0.105	0.219
		(0.600)	(0.430)		(1.30)	(1.32)
South region		-0.041	-0.070		0.060	0.293
		(-1.58)	(-0.880)		(0.780)	(14.63)
West region		0.180	0.137		0.247	0.465
		(6.05)	(1.56)		(3.01)	(4.91)
Food Stamps		-0.002	0.058		0.094	0.042
		(-0.070)	(1.38)		(0.840)	(1.63)
Pseudo R-squared	0.110	0.142	---	0.07	0.100	---
Breusch-Pagan LM test for random effects (test stat and p-value)			386.02 (0.000)			92.29 (0.000)
Hausman specification test (test stat and p-value)			41.95 (0.000)			39.98 (0.0005)

Breusch-Pagan test for heteroscedasticity (test stat and p-value)			---			669.68 (0.000)
N families			874			349
N children	3397	3397	1941	1619	1619	768

Note: Huber t-statistics (standard models) and t-statistics (fixed effects models) in parentheses, and intercept not shown.