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TEENAGE MOTHERS, LABOR FORCE PARTICIPATION, AND WAGE RATES

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Résumé — Dans cette étude nous avons ébauché et évalué les coefficients d'un modèle de la participation de la population active et les taux de salaire proposés par Heckman. Ce modèle-ci est attirant parce qu'on n'a pas supposé que les femmes qui travaillent constituent un échantillon aléatoire de toutes les femmes. On peut montrer qu'une telle supposition les cas extrêment spéciaux exceptés, mène aux estimations biaisées des paramètres des équations qui déterminent les salaires et les heures de travail. Nos résultats indiquent que l'effet indépendant de l'âge à la première naissance sur le salaire de marché, une fois que les autres variables sociales et démographiques sont contrôlées, est négligeable; l'éducation et l'expérience sont les déterminants importants du salaire de marché. L'âge à la première naissance n'a pas d'impact sur le salaire du réserve, même quand l'éducation et d'autres mesures en fécondité sont maintenues constantes. Cet impact est significatif, grand et positif que pour les blancs; augmentant leur âge à la première naissance, diminue leur disposition au travail.

Abstract — We have outlined and estimated the coefficients of a model of labour force participation and wage rates proposed by Heckman. This model is attractive because it does not assume that working women constitute a random sample of all women. Such an assumption, except in highly special cases, can be shown to lead to biased estimates of the parameters of the equations which determine wages and hours of work. Results indicate that the independent effect of age at first birth on the market wage, once other social and demographic variables are controlled, is negligible; education and experience are the important determinants of the market wage. Age at first birth does have an impact on the reservation wage, even when education and other fertility measures are held constant. This impact is significant, large, and positive only for whites; increasing their age at first birth lowers their propensity to work.

Key Words — **wage rates, labour force participation rates, teenage childbearing**

During the past few years, considerable attention has been focused on the subsequent effects of early childbearing. This attention has been prompted by large numbers of births to teenaged mothers, coupled with unchanging birth rates among women under 15 years of age and slowly decreasing rates among women age 15-19. Yet, a recent conference concluded that remarkably little is known about the social, economic, or demographic impact of early childbearing on the parents, children or community (Chilman, 1980).

Perhaps the most serious concern involves the economic consequences of early childbearing. Until very recently, research on this topic was virtually nonexistent (Trussell, 1976). Coombs and Freedman (1970) did examine the relationship between premarital pregnancy and later economic achievement. They found that couples with a premarital pregnancy suffered a substantial disadvantage when compared with other couples. Their income differential disappeared only if education and marriage duration were controlled. Moreover, with such controls, their unfavourable asset position was only reduced, not eliminated. Several recent studies have directly examined the impact of age at first birth on wage rates, labour force participation rates, and income of women (McLaughlin, 1977; Hofferth and Moore, 1979; Hofferth, Moore and Caldwell, 1978). Although these studies contain results with important policy implications, they are all based on ordinary least squares (OLS) regression analysis, and some are based on samples of working women only. Labor economists have shown, however, that such methodology is inappropriate, but their results have not reached the wider audience of demographers and sociologists. In this paper, therefore, we have two goals. First, we examine in detail why estimates based on samples of only working women, particularly those derived from OLS, lead to biased and misleading results and employ a methodology which we hope will be more widely adopted by demographers interested in the effects of demographic variables on labour force participation and wage rates. Second, we assess directly the effect of age at first birth on these two economic variables. Our primary concern is with this substantive issue. We cannot evaluate fully the overall consequences of teenage pregnancy, but we hope to provide a first step by discovering its effect on the labour force participation and wage rates of the mother when she is over age 25.

The Model

There are two behavioral equations which determine the probability that a woman works, the hours which she works, and her observed market wage. The first is the traditional equation expressing the market wage W, also called the offered wage, as a function of certain variables assumed or known to affect the wage:

$$W_i = f(X_i) \tag{1}$$

where W_i is the wage that the ith woman could command in the market and X_i is a vector of characteristics of the woman. The second equation determines the shadow wage W^*, also called the asking or reservation wage, for the woman's non-market time and may be written as

$$W_i^* = g(h,Z_i) \tag{2}$$

where h is the hours of work and Z_i is a vector of characteristics assumed to affect the value of non-market time. This particular model was described by Heckman (1974). His paper contains a formal derivation of equation (2).

If, when deciding whether or not to work, $W_i^* > W_i$, then the value of the woman's non-market time exceeds the wage she could obtain in the market, and she will not work. If, $W_i > W_i^*$, and a woman is free to vary the hours that she does work, then her hours will be increased up to the point where $W_i^* = W_i$, when the marginal gain from working an extra hour just equals the marginal value of an hour otherwise occupied. Here, the implicit assumption is made that the partial derivative of W^* with respect to h is positive. Otherwise the system would have no equilibrium solution. It is certainly plausible that the marginal value of a woman's non-market time increases as time spent on household duties decreases; nevertheless, this assumption is explicitly tested below. Following Heckman, the particular case in which a woman works all available hours and still $W > W^*$ is dismissed as "empirically uninteresting". If a woman works, $W_i^* = W_i$, while if she does not work, then $h_i = 0$; hence, the basic constraint $h_i(W_i - W_i^*) = 0$ holds for all women free to choose their working hours. If a woman does work, then equations (1) and (2) recursively determine the hours of work; hours adjust until the asking wage W_i^* rises to meet the market wage W_i, which is assumed to be independent of hours worked. The assumption that women are indeed free to alter their working hours until the shadow wage equals the market wage is one which should arouse some skepticism. Normally, the conditions of employment, including hours of work, are dictated by the employer. A woman may better be able to vary the number of weeks she works. In the empirical section below, we use both variants of the measure of market time. Our particular choice of sample, discussed below, also renders the assumption of freely chosen market time more plausible. The model ignores unemployment, since unemployment constrains women to work zero hours even though in this situation, $W_i > W_i^*$.

The determinants of the market wage are fairly well understood. Education and the length of work experience, on the job training, have been found to affect positively the market wage (Mincer, 1962; Mincer and Polachek, 1974). The characteristics which determine the asking wage have not been fully explored. The vast literature on the determinants of labour supply (*e.g.* Bowen and Finegan, 1969; Sweet, 1973; and Cain, 1966) suggests a number of variables which might affect the shadow wage. Our purpose here is to determine whether other variables, including age and, most importantly, age at first birth affect the market and shadow wages. Our maintained hypothesis is that age at first birth has no significant impact on the offered wage but may affect the asking wage. Once education and the length of work experience have been controlled, we could think of no reason why age at first birth should have any effect on the market wage; whereas an employer would be vitally interested in a woman's market skills, he would have no reason to care about or perhaps even no opportunity to know her age at the time of her first birth. On the other hand, we reasoned that age at first birth might well affect the woman's shadow or reservation wage, perhaps because the options in life for a teenage mother are constrained. A birth colors perceptions long after the event; its timing in the life cycle might be expected to affect the perceived opportunity cost of market time. Although we have no strong prior convictions about the direction of the relationship, we feel that the above considerations might imply that a younger age at first birth, *ceteris paribus,* would be associated with a lower value of non-market time.

In order to estimate equations (1) and (2), the functional forms and the underlying stochastic structure must be specified. Following Heckman, we assume that there exists a suitable monotonic transformation of the dependent variables so that both may be expressed as linear functions of their arguments. If $l(\,)$ is such a transformation, then equations (1) and (2) may be re-expressed as

$$l(W_i) = X_i'b + \mu_i \tag{3}$$

$$l(W_i^*) = h_i\gamma + Z_i'\beta + \epsilon_i \tag{4}$$

where b and β are vectors of coefficients in the market and asked wage equations respectively, γ is the coefficient of hours of work, and μ_i and ϵ_i are the associated random errors. The errors are assumed to arise from the inability of the researcher to specify the variables X_i and Z_i fully so that they represent unobserved (and unobservable) characteristics known to the women but unknown to the investigator. For estimation purposes, we assume that ϵ and μ are jointly normally distributed with means zero, variances σ_ϵ^2 and σ_μ^2, and covariance $\sigma_{\epsilon\mu}$.

The class of monotonic transformations considered here is limited to the class of power transformations suggested by Box and Cox (1964):

$$l(W) = \frac{W^\lambda - 1}{\lambda} \tag{5}$$

This class of transformations is extremely flexible and includes the log($\lambda = 0$) and linear ($\lambda = 1$) transformations as special cases. Since the parameter λ is estimated as well, the appropriate transformation may be ascertained.

If a woman works, then at the hypothetical, unobservable situation of zero hours of work, $l(W_i^*) < l(W_i)$ or,

$$Z_i'\beta - X_i'b < \mu_i - \epsilon_i \tag{6}$$

In such a case, hours of work h rise until $l(W) = l(W^*)$. *Given that condition (6) holds,* the reduced form for wages and hours becomes

$$l(W_i) = X_i'b + \mu_i \tag{7}$$

$$h_i = \frac{1}{\gamma}\ (X_i'b - Z_i'\beta) + \frac{\mu_i - \epsilon_i}{\gamma} \tag{8}$$

Equations (7) and (8) are ordinary wage and labour supply functions; for example, if $l(W)$ were linear, then $\frac{1}{\gamma}$ would be the uncompensated response of labour supply of working women to a unit change in the offered wage.[1]

It is tempting to stop at this point and estimate equations (7) and (8) on a sample of working women directly by OLS. Such a step, however, ignores the crucial feature of this model. Equations (7) and (8) hold only if condition (6) holds. Hence the distributions of the stochastic components of (7) and (8) are *conditional* distributions which depend upon the observed characteristics of the ith woman. But these same characteristics appear as explanatory variables in equations (7) and (8). Thus the errors in these equations are correlated with the explanatory variables. As a result, OLS will not yield unbiased estimates. Furthermore, this problem will not disappear as the number of observations increases[2] so that the OLS estimates will not even be consistent. The problem arises not because of deficiencies in the OLS estimating technique itself, but because the sample of working women is not selected randomly from all women since the decision to work is in part determined by individual characteristics. No method of directly estimating equations (7) and (8) can avoid this selection bias, and interpretation of the parameters so estimated is difficult, if not impossible, since meaningful structural parameters are inextricably mixed with the parameters which determine whether a woman enters into the sample. For example, determinants of the propensity to work, such as the presence of children (Hofferth, Moore and Caldwell, 1978) can attain spurious significance in the wage equation, when, in fact, such variables do not affect the wage (Heckman, 1980).

This point is a difficult one, and closer examination may help to illuminate the trouble. The particular adjustment in hours needed to equate $l(W^*)$ with $l(W)$ depends on the magnitude of the discrepancy between the quantities $Z_i'\beta - X_i'b$ and $\mu_i - \epsilon_i$. Since this discrepancy must always be negative in order for a woman to enter the sample of working women, the error term $\mu_i - \epsilon_i$ cannot be independent of the explanatory variables. Independence is lost when the sample is confined to working women only. More explicitly, since neither the wage nor hours supplied will be observed unless $Z_i'\beta - X_i'b < \mu_i - \epsilon_i$, then the expected values of hours and wages are conditional expectations:

$$E[l(W_i) | \mu_i - \epsilon_i > Z_i'\beta - X_i'b] = X_i'b + E[\mu_i | \mu_i - \epsilon_i > Z_i'\beta - X_i'b] \tag{9}$$

and

$$E[h_i | \mu_i - \epsilon_i > Z_i'\beta - X_i'b] = \frac{1}{\gamma}\ (X_i'b - Z_i'\beta) + \frac{1}{\gamma}\ E[\mu_i - \epsilon_i | \mu_i - \epsilon > Z_i'\beta - X_i'b] \tag{10}$$

When expressed this way, it can be clearly seen that regressions, such as those reported by Hofferth, Moore and Caldwell (1978), which include only the first terms on the right-hand sides of equations (9) and (10), estimating directly equations (7) and (8), suffer from the familiar problem of omitted variables. Unless the conditional expectations of these error terms are zero, the coefficients estimated from equations (7) and (8) will in general be biased and corresponding tests of their significance invalid. It can also be seen that if the random variables μ_i and $\mu_i - \epsilon_i$ are uncorrelated, or independent if the normality assumption is not invoked, then direct estimates of (7) will provide unbiased estimates of the parameters b. However, these estimates will not be efficient since the cross-equation restriction that the coefficients of X in the two equations should be proportional, with proportionality factor $1/\gamma$, has not been imposed. Direct estimation of (8) on a sample of working women can yield unbiased estimates only if the variance of $\mu_i - \epsilon_i$ is zero; the equation is an identity with no random component. Proofs of the conditions necessary for the conditional expectations of the error terms to vanish, may be found in the appendix of this paper. These observations suggest tests of whether sample selection bias is important; these tests will be discussed more fully below.

An examination of the likelihood function shows directly that parameters estimated only from working women must in general be biased. If a sample of n women contains k women who work and n-k women who do not, then the likelihood function may be written as

$$L = \prod_{i=1}^{k} j(h_i, W_i) \times \prod_{i=k+1}^{n} Pr[W_i < W_i^*]_{h=0} \qquad (11)$$

where $j(h_i, W_i)$ is the joint density function of hours and wages of women who work and $Pr(W_i < W_i^*)_{h=0}$ is the probability that at zero hours of work the offered wage falls short of the asking wage. The particular functional forms of $j(h_i, W_i)$ and $Pr[W_i < W_i^*]_{h=0}$ are derived in the appendix to Heckman's (1974) article.[3] Since the parameters to be estimated enter into both parts of the likelihood function, estimation based on only part of it will yield biased results. Maximizing the likelihood function (11) with respect to all the parameters of the model, including σ_ϵ^2, ϵ_μ^2, $\sigma_{\epsilon\mu}$, and λ, yields consistent and efficient parameter estimates which are asymptotically normally distributed with means equal to the true parameter values. It is helpful to redefine the parameters of the model by letting $\rho_i = \mu_i - \epsilon_i$. Then, the conditional expectation of the error term in equation (9) will not vanish unless the covariance between ρ and μ is zero and the conditional expectation of the error term in equation (10) will not vanish unless the variance of ρ is zero. Thus, one has a simple test of whether selection bias is important in each equation.

Note, finally, that this model is much more attractive than the alternative of estimating equation (8) for *all* women by assigning zero hours (McLaughlin, 1977; Hofferth and Moore, 1979) to women who do not work. Such a procedure will invariably produce biased results since the regression line is twisted by the many women who do not work (Goldberger, 1964:252). The intuitive argument here is that for some women, non-market time is so much more valuable than their potential market wage, that they would ideally like to work negative hours; they would like to buy additional hours of non-market time, but, of course, are constrained from working fewer than zero hours. The Heckman procedure is also more attractive than estimating a composite income (= hours × wage) equation (Hofferth and Moore, 1979), since the separate effects of independent variables on the two components can be disentangled. It should be noted that even if one wanted to estimate an income equation, the procedure employed by Hofferth and Moore of assigning zero income to women who do not work and estimating the equation by OLS will produce biased results since the dependent variable is bounded. The traditional single equation model used in such cases was developed by Tobin (1958). The present model devised by Heckman is an extension of the Tobit regression to allow the simultaneous estimation of parameters common to both equations.

The most common objection to the Heckman model is on grounds of convenience and cost, not theory. We found, however, that the likelihood function was not hard to code and that the cost of estimating the model averaged about $15. Optimization of the likelihood function (11) requires numerical techniques. The particular technique employed here was developed by Berndt, *et al.* (1976). To ensure that a unique maximum had been reached, we altered the starting values. In all cases, the same maximum likelihood estimates were obtained. Heckman (1976, 1979, 1980) has suggested another technique which explicitly accounts for the sample selection bias and provides consistent estimates of the parameters. Specifically, initial estimates of β/σ_p and b/σ_p are obtained by running a probit regression in which the dependent variable is 1 if the woman works and 0 if she does not. These parameter estimates, in turn, provide estimates of a variable, the λ_i derived in the appendix, corresponding to the conditional expectations of the error terms in equations (7) and (8). When this variable is explicitly included as a regressor, each equation can be estimated using OLS. It should be noted that the resulting parameter estimates are not efficient since the cross-equation restrictions are ignored. Other investigators may find his alternative approach more appealing and will undoubtedly find that the discussion of sample selection bias and informal tests for such bias is both useful and excellent.

The Data

The analysis is based upon Cycle I of the National Survey of Family Growth (NSFG) conducted by the National Center for Health Statistics. The NSFG was designed to provide information about fertility, family planning intentions and activity, and other aspects of maternal and child health which are closely related to childbearing. Data on each of these topics were collected in personal interviews with approximately 9800 women aged 15-44 who either had ever been married or had children of their own living in the household. Interviews were conducted between July 1973 and February 1974, and centered on September 13, 1973. Respondents were selected by a multi-stage, area probability, cross-sectional sample of households in the continental U.S.A.

This particular data set was chosen because it contains detailed information on the fertility experience of women. Other data sets, such as the National Longitudinal Survey, probably contain more detailed economic information, but at the expense of less adequate fertility data. The sample was limited to married women between the ages of 25 and 44 who had delivered at least one child which was between the ages of 12 and 30 at the time of the sample. Women below the age of 25 were excluded from the sample because for a sizeable proportion of these, additional schooling remained an alternative. The assumption that women are able to alter their hours or weeks of work until the marginal values of market and non-market time become equal is more plausible among this set of older women than, for example, among teenagers. In addition, since women entered the NSFG sample only by marriage or motherhood, the younger women are non-representative of all women in their birth cohorts. Finally, because the influence of age at first birth (AFB) on labour force participation and wage rate is the subject of our concern, we restricted our attention to parous women. After eliminating 847 women, 17 per cent of all respondents, because of problems in calculating wages or hours worked or because of inconsistencies between the respondent's income, wages, and hours worked information, a total of 4183 respondents remained. We then split the sample into three groups: one black sample of 1186 women; a first white sample of 1533 women, used for estimation; and a second white sample of 1474 women, used for testing. White respondents were assigned randomly to the two samples.

Attached to each respondent is a weight designed to indicate the number of married women in the U.S. whom the sample woman represents. It would, therefore, appear necessary to account for these weights explicitly in the estimation scheme. It is reassuring to note that, in this model, the estimates based on the weighted and unweighted samples are nearly identical.[4]

There is no unambiguously preferred choice for the unit in which to measure labour supply. Perhaps the most obvious choices are hours worked per year or weeks worked per year. It could be argued that women are freer to adjust the number of weeks than the number of hours per week which they work. On the other hand, weeks are rather lumpy units, and considerable variation among women is eliminated by this choice. Because of genuine uncertainty in this regard, separate sets of coefficients for weeks and hours were developed as the units of measurement of labour supply in equation (11); results for each set are presented below. The particular choice of a year as the time period of analysis was dictated by the data.

Education and AFB were entered as dummy variables rather than as cardinal variables so that any non-linear effects could be captured. Three dummy variables for education, indicating a high school diploma, some college, and at least a college degree were employed. The three dummy variables for AFB denote entry into motherhood at ages 18-19, ages 20-24, and ages 25-29. The constant term in the regression absorbed the effect of both AFB under 18 and education not extending through high school. A separate dummy indicated whether a woman had completed at least one year of special schooling, including business, technical, or nursing school. The education and AFB dummy variables were entered into both equations. The number of children under 6 years of age present in the household and the number of children aged 6 to 18 present in the household were entered separately in the shadow wage equation so that any differential effects could be noted. In addition, a dummy variable indicating the presence of any children under 6 in the household was entered to test whether the mere presence or the actual number of children had a significant effect on the reservation wage. The total number of children in the household was entered into the market wage equation in order to test the significant effect reported by Hofferth, Moore, and Caldwell (1978). Our expectation was that this variable would not affect the market wage, if work experience is controlled (Mincer and Polachek, 1974). Work experience, which has been shown to affect positively the market wage in numerous studies, was measured as the accumulated number of years of actual work experience.[5] All income other than that earned by the woman was entered in the shadow wage equation as the measure of outside income available to the family; because there was no way to identify separately any welfare or unemployment benefits which may be conditional on the labour supply of the wife, such income would be included in this measure. One might expect the age of a woman, *ceteris paribus,* to raise the asking wage. We had no clear expectations about the effect of age on the market wage. Holding years of schooling and experience constant, an older woman has more non-job experience, and thus may have accumulated more wisdom than her younger colleagues. On the other hand, persons who have received the same amount of education and have the same length of work experience at a younger age may be perceived as more able (Boulier and Pineda, 1975). We included age in both equations to determine empirically its role. Finally, whether or not the first birth is conceived within marriage, premaritally conceived and later legitimized, or illegitimate, might be posited to have an effect on the shadow wage even after AFB is controlled; weak and strong effects, respectively, have been noted in examinations of the pace of childbearing (Trussell and Menken, 1978) and marital dissolution (McCarthy and Menken, 1979). Since we are particularly interested in this question, two dummy variables indicating an illegitimate and legitimized first birth were included in the shadow wage equation.

Empirical Results

We originally hypothesized that with education included in the market wage equation, AFB would prove to have no significant effect. Based upon asymptotic likelihood ratio tests we confirmed that the effect of AFB in the market wage equation was insignificant both for blacks and whites and for either the hourly or weekly unit of measurement of labour supply. A similar test confirmed that the number of children was an insignificant determinant of the hourly wage. This result, which conforms to economic theory, contrasts with the finding of Hofferth, Moore, and Caldwell (1978). Since their results were based on a sample of working women, however, the structural parameters of the wage equation cannot be untangled from the parameters which determine whether a woman works. These fertility variables were then dropped from the market wage equation and the remaining parameters re-estimated. Two dummy variables indicating whether the first birth was illegitimate, or premaritally conceived but later legitimized, were originally included in the shadow wage equation. The null hypothesis that the coefficients of these variables were equal could not be rejected at even the 0.2 level, so these two variables were combined into a single dummy variable which indicates that the first birth was not conceived within marriage. Final estimates are presented in Table 1.

Previous empirical studies (*e.g.*, Mincer and Polachek, 1974; Cain, 1966; Bowen and Finegan, 1969; Sweet, 1973) found that the coefficients for blacks and whites were different. To test this hypothesis, we estimated the coefficients for blacks and whites combined and compared the estimated value of the likelihood function with the product of the values obtained from estimating each race separately. Based on asymptotic likelihood ratio tests, the hypothesis that the two sets of coefficients were the same was rejected conclusively. This result contrasts starkly with that of Hofferth, Moore, and Caldwell (1978), who estimated the two equations separately on a sample of working women. They found that coefficients for blacks and whites in either the wage or hours equation were insignificantly different. It should be emphasized, however, that since their results were estimated for samples of working women only, the coefficients are biased and resulting tests of hypotheses invalid.

Close examination of the results presented in Table 1 reveals several interesting findings. First, the variance of $\rho = \mu - \epsilon$ is significantly greater than zero. Thus, it is inappropriate to estimate the labour supply equation (8) directly from the sample of working women. Second, the covariance between ρ and μ is significantly different from zero for blacks — though not strongly so when labour supply is measured in hours — but at best marginally different from zero among whites. Thus, there is no strong evidence that the coefficients of the wage equation (7) for whites could not be directly estimated on the sample of working women. Both of these findings are consistent with those of Heckman (1980), who estimated the model on a sample of 1735 white women, who were married with spouse present and with husbands working in the previous year (1966), taken from the 1967 National Longitudinal Survey of Work Experience of Women Aged 30-44. The value of λ is significantly above zero (~ 0.16) for whites but not significantly different from zero for blacks using either measure of labour supply. Thus, while the logarithmic transformation of wages is appropriate for blacks, it is not the correct transformation for

TABLE 1

	Parameter	Description of Parameter	Hours: White Coefficient	Hours: White |Mean/se|	Hours: Black Coefficient	Hours: Black |Mean/se|	Weeks: White Coefficient	Weeks: White |Mean/se|	Weeks: Black Coefficient	Weeks: Black |Mean/se|
VARIANCE-COVARIANCE	σ_ρ^2	Variance of μ-ϵ (= ρ)	.05724	2.51	.05167	3.05	.09228	2.99	.08798	3.93
VARIANCE-COVARIANCE	$\sigma_{\rho\mu}$	Co-variance of μ and μ-ϵ	-.00253	0.31	.01270	1.83	.02830	1.83	.04609	3.83
VARIANCE-COVARIANCE	σ_μ^2	Variance of μ	.30846	20.07	.16150	16.54	.31638	18.49	.17736	14.17
SHADOW WAGE	γ	Hours (hundreds) or weeks	.02014	4.92	.02113	6.06	.00949	5.80	.01085	7.72
SHADOW WAGE	β_0	Constant	.20339	1.09	.45109	3.56	.02428	0.13	.31155	2.52
SHADOW WAGE	β_1	Other income (thousands of dollars)	.00389	3.01	-.00017	0.12	.00473	3.22	.00176	1.05
SHADOW WAGE: EDUCATION	β_2	Education = H.S. diploma	.17664	3.08	.20598	5.06	.16274	2.90	.21107	5.30
SHADOW WAGE: EDUCATION	β_3	Education = some college	.40526	4.90	.30360	5.89	.37303	4.64	.29738	5.96
SHADOW WAGE: EDUCATION	β_4	Education > college degree	.70489	8.69	.88320	12.15	.67940	8.42	.88618	12.63
SHADOW WAGE: EDUCATION	β_5	Special education	.02259	0.36	.22339	5.41	.03644	0.60	.22346	5.61
SHADOW WAGE: FERTILITY	β_6	18 ≤ AFB < 20	.00301	0.12	-.05131	2.28	-.01673	0.49	-.07326	2.74
SHADOW WAGE: FERTILITY	β_7	20 ≤ AFB < 25	.07712	2.66	-.03023	1.51	.05771	1.71	-.04221	1.72
SHADOW WAGE: FERTILITY	β_8	25 ≤ AFB < 30	.18059	3.84	.04732	1.62	.18345	3.84	.04535	1.30
SHADOW WAGE: FERTILITY	β_9	Presence of children < 6	.04811	1.50	-.03839	1.42	.07502	1.90	-.05529	1.69
SHADOW WAGE: FERTILITY	β_{10}	# children < 6	.09225	3.38	.10322	4.51	.11176	3.61	.13046	5.31
SHADOW WAGE: FERTILITY	β_{11}	# children aged 6-18	-.00682	1.14	.00092	0.21	-.01255	1.65	.00145	0.27
SHADOW WAGE: FERTILITY	β_{12}	First child conceived out of wedlock	.03299	1.63	.05354	2.38	.04794	1.85	.08852	3.14
SHADOW WAGE	β_{13}	Age	.01153	2.62	-.00013	0.04	.01338	3.01	-.00073	0.26
MARKET WAGE	b_0	Constant	.49141	3.05	.74457	6.94	.38008	2.30	.68055	6.18
MARKET WAGE: EXPERIENCE	b_1	Years of work	.04838	4.96	.04788	6.20	.06313	5.72	.06129	7.61
MARKET WAGE: EXPERIENCE	b_2	Years of work squared	-.00109	3.91	-.00132	5.65	-.00150	4.36	-.00167	6.63
MARKET WAGE: EDUCATION	b_3	Education = H.S. diploma	.18286	3.39	.24308	6.54	.17376	3.19	.25749	6.73
MARKET WAGE: EDUCATION	b_4	Education = some college	.44284	5.60	.37018	8.08	.41249	5.22	.38947	8.14
MARKET WAGE: EDUCATION	b_5	Education ≥ college degree	.79975	11.31	1.01202	17.27	.80646	11.23	1.06551	16.26
MARKET WAGE: EDUCATION	b_6	Special education	.07437	1.29	.22663	5.85	.09375	1.61	.23435	5.86
MARKET WAGE	b_7	Age	.00166	0.37	-.01104	3.77	.00056	0.13	-.01365	4.51
	λ		.16398	9.26	-.00264	0.09	.16730	9.80	-.00223	.08

whites. This result is not consistent with Heckman's (1974) finding that $\lambda \sim 0$ for whites; in subsequent articles he has assumed that the logarithmic transformation is appropriate. The coefficient of either hours or weeks is positive and significantly above zero for both blacks and whites, respectively; the values of γ are, moreover, similar for each race. The fact that the values of γ are similiar but the values of λ are different indicates a racial difference in response of labour supply to an exogenous increase in the wage, however. An exogenous increase in the wage is indicated by a change in the constant b_0 in the offered wage equation. From equation (8), we see that the response of labour supply of working women can be estimated as $\partial h/\partial b_0 = 1/\gamma$. A change of b_0 by one unit would raise the labour supply of blacks by 92 weeks and that of whites by 105 weeks. These responses seem excessive unless one realizes that b_0 is measured in units of $l(W)$ dollars; hence a change in b_0 of one unit raises the market wage of whites and blacks by 123 per cent and 173 per cent, respectively. Thus a 10 per cent increase in the wage would elicit 8.5 and 5.3 additional weeks of work, respectively, among white and black women. While the elasticities[6] implied by this response (3.25 and 2.14, respectively) are high, they are consistent with more recent findings for women; indeed, Heckman (1980) has argued that elasticities estimated from traditional labour supply functions are too low. It should be noted that the current thinking among labour economists is that previous elasticities of female labour supply estimated by OLS were far too low. When the exercise is repeated using the results from the equations based on hours, the elasticities are higher (4.50 and 2.93, respectively) but still consistent with current thinking.

Turning first to the examination of the effect of fertility variables on the asking wage, we see that AFB does not have a monotonic[7] impact for blacks; among whites this impact is monotonic only in the hours equation. For whites, delaying a first birth until ages 18-19 does not significantly alter the reservation wage. After this age group, postponing entry into motherhood raises the reservation wage significantly. For blacks the situation is more confusing. Delaying childbearing until 18-19 lowers the reservation wage significantly, but delaying motherhood until ages 20-24 or 25-29 has only a marginally significant positive effect. Since all the dummy variables measure an incremental effect relative to the constant term (in this case AFB < 18), the negative effect for an AFB of 18-19 would be explained by any circumstance which lowered, relative to other AFB groups, the propensity to work among women whose first birth occurred before age 18. One hypothesis, though not testable with these data, is that young black women at the time of their first birth have fewer alternatives to accepting welfare assistance than older black women, and that this acceptance persists in later life. That it is the under-18 group which is the anomaly can perhaps be more easily demonstrated by the observation that if the constant term had absorbed the effect of an AFB of 18-19, the incremental effect for the next two age groups would have been positive and the incremental effect for the group under 18 anomalously positive. The mere presence of children under six years of age is significant only in the weeks equation; however, the effect is in opposite directions by race. The reservation wage is raised among whites and lowered among blacks by the presence of children. The implication is that the presence of children under six years of age is associated with a higher propensity to work among blacks and a lower propensity to work among whites. The number of children under six has a significantly positive effect on the asking wage for both races, but the number of children aged 6-18 has no significant impact. The obvious interpretation of this finding is that increasing the number of children who potentially would be at home with the mother increases the reservation wage; when these children are old enough to be at school, they have no significant effect on the shadow wage and hence on the propensity to work. These effects of children on labour force participation are consistent with the results of previous studies (*e.g.*, Sweet, 1973; Cain, 1966; Bowen and Finegan, 1969). Other income has a significant effect on the asking wage only for whites. Increasing other income by \$1000 would decrease their supply of labour by 19.3 hours or 0.498 weeks; the implied elasticities are -0.414 and -0.362 for the hours and weeks specifications, respectively. Increasing educational attainment has a monotonic, significant positive impact on the reservation wage for both races, whereas the impact of special education is significantly positive only for blacks. This racial difference in the significance of special education may reflect recent findings (Mott and Shaw, 1978) that black females seek special training for job-related reasons, while white females participate as a means of continuing education and that a higher fraction of whites fails to complete training programs. Turning to the market wage equation we see that years of work and years of schooling have significant positive effects on the market wages. Completing a year of special education has a significant positive effect only for blacks. Hence the same racial difference in the significance of special education is found in the market and shadow wage equations; in the market wage equation, this difference may simply be a consequence of the fact, reported above, that a black woman with some special education is more likely to have entered the program for job-related reasons and more likely to have completed the training. The square of number of years of work experience is significantly negative for both races, indicating that although additional work experience raises the wage, it does so at a decreasing rate. Finally, increasing age lowers the market wage among blacks, but has no effect among whites, once other variables are controlled.

The estimated coefficients can be used to evaluate the impact of a change in one variable, holding the remainder constant at their mean values, on the probability of working. As Table 2 indicates, increasing educational attainment monotonically decreases the probability of not working. This qualitative effect could have been inferred from Table 1, since the values of the coefficients of each education variable are higher in the market wage equation than in the asking wage equation. Since AFB enters only the reservation

TABLE 2 PROBABILITY OF NOT WORKING AS AGE AT FIRST BIRTH AND EDUCATION CHANGE.*

	Education Only Changes				Age At First Birth Only Changes			
	<HS	HS	Some College	≥College	<18	18-19	20-24	25-29
Black (hours)	.288	.235	.197	.130	.272	.203	.230	.345
Black (weeks)	.269	.220	.177	.111	.258	.185	.215	.310
White (hours)	.524	.514	.462	.368	.376	.380	.502	.669
White (weeks)	.497	.483	.446	.336	.390	.369	.465	.627

* All other variables held constant at their mean values.

TABLE 3 OBSERVED AND PREDICTED LABOR FORCE PARTICIPATION.

	Predicted LFP			
	Parameters Based On			
Observed LFP	Hours Worked		Weeks Worked	
	Yes	No	Yes	No
Yes	505	182	529	158
No	222	565	257	530

wage equation, then, qualitatively, the effect of changing age at entry into motherhood on the probability of working can be directly inferred; increasing AFB raises the perceived value of non-market time and thereby lowers the probability of working. Quantitative results are also presented in Table 2.

The estimated parameters can also be used to test to see how well the model predicts. Specifically, we employed the parameter estimates from the first white file to predict whether women in the second white file would be workers or non-workers. The results, shown in Table 3, are gratifying. Seventy-three and 72 per cent of the women were correctly classified using the hours and weeks equations, respectively. Since the labour force participation rate of women in the first sample is less than 0.5, the best prediction one can make in the absence of a model is that no woman works. With this rule, 53 per cent of the predictions in the second sample would be correct. Hence, the model increased the predictive power by approximately 36 per cent.

Conclusions

In summary, we have outlined and estimated the coefficients of a model of labour force participation and wage rates proposed by Heckman (1974). This model is attractive because the assumption is not made that working women constitute a random sample of all women. Such an assumption, except in highly special cases, can be shown to lead to biased estimates of the parameters of the equations which determine wages and hours of work.

Our results indicate that the independent effect of age at first birth on the market wage, once other social and demographic variables are controlled, is negligible. This result was exactly as we had hypothesized, since employers are not likely either to know or care about the age of a woman at the time of her first birth; education and experience are the important determinants of the market wage. On the other hand, age at first birth does have an impact on the reservation wage, even when education and other fertility measures are held constant. This impact is significant, large, and positive only for whites. Among whites, increasing the age at first birth lowers their propensity to work; increasing the age at first birth from under 18 to 18-19 has little influence, but increasing it from 18-19 to 20-24 or 25-29 decreases the proportion of women who work by approximately 20 per cent and 50 per cent, respectively.

It is, however, very possible that age at first birth affects educational attainment, which we, as others, have shown to have a profound positive effect on both the market wage and labour force participation. While neither the NSFG data set, nor any other that we know, is sufficient to address conclusively the problem of causality between entry into motherhood and educational attainment,[8] we can give some notion of the impact by returning to the model. Suppose that age at first birth determined educational attainment, and that this effect could be captured by the observed educational distributions of the four age-at-first-birth categories. These distributions, together with the market wage which they imply when all other wage determinants are held constant at their mean values, are shown in Table 4. It can be clearly seen that if the different age-at-first-birth categories did imply the corresponding distributions of educational attainment, then age at first birth would have a profound impact on the market wage, albeit only indirectly. Postponing motherhood until the age of 20 would raise the wage by approximately 20 per cent when compared with an age at first birth below 18. Hence we are led to the conclusion suggested by Trussell (1974) that if teenage motherhood does have any significant economic impacts on the woman, such effects are likely to operate through education.

The results reported above could obviously be obtained by merely altering the educa-

TABLE 4 EDUCATION DISTRIBUTIONS OF AGE-AT-FIRST-BIRTH CATEGORIES.

	Age At First Birth							
	White				Black			
Education	<18	18-19	20-24	25-29	<18	18-19	20-24	25-29
<HS	.804	.354	.156	.113	.711	.417	.248	.271
HS	.184	.560	.562	.396	.222	.441	.539	.372
Some College	.013	.073	.176	.208	.061	.122	.164	.163
≥ College	.000	.013	.106	.283	.006	.020	.049	.194
Market Wage *								
Hours equation	2.18	2.40	2.66	2.95	2.10	2.29	2.46	2.73
Weeks equation	2.03	2.22	2.46	2.74	1.94	2.14	2.30	2.57

* All other variables held constant at their mean values.

tional distribution of a *given* age-at-first-birth category. Thus, successful efforts to encourage teenage mothers to stay in school or return to school would appear to have a positive impact on their market potential. One obvious reason why many women have had their education truncated by a teenage pregnancy is that they were forced to leave school. Such forced exclusion from public schools of teenagers who are mothers or who are married or pregnant is now illegal under Title IX of the Educational Amendments of 1972, effective July 12, 1975. Therefore, one obstacle to educational achievement has been eliminated. It is still not known, however, what proportion of such women will continue to attend school.

Admittedly, our focus has been rather narrow. The overall economic well-being of mothers and their children certainly depends on more than their own wage rates and propensities to work. If, as the work of Coombs and Freedman (1970) suggests, a younger age at first birth tends to truncate the education of husbands of women who marry, then family economic attainment would be lessened as well. A younger age at first birth has been found to be strongly associated with a more rapid pace of subsequent childbearing (Bumpass, Rindfuss, and Janosik, 1978; Trussell and Menken, 1978) and is thought to be associated with a more rapid rate of marital dissolution (McCarthy and Menken, 1979). Other things being equal, these two findings further suggest that the trajectory of economic advancement might be lowered by young motherhood.

Acknowledgment

We have followed Jim Heckman's discussions of his model closely; his comments are greatly appreciated. Jane Menken and Bryan Boulier provided detailed suggestions which we gratefully incorporated.

Footnotes

1 The interpretation of this effect is somewhat tricky; $1/\gamma$ is the effect on the hours worked *of working women* with respect to a unit change in $l(W)$. As will be discovered in the text, such a change would increase the sample of working women since the market wage would be higher than the unchanged reservation wage at the hypothetical situation of zero hours of work for a higher proportion of women. The effect on the hours of work offered by all women who would be working after the exogenous shock is given by the derivative of equation (10) with respect to b_0. The extra term $\partial E[h|condition\ (6)]/\partial b_0$ accounts for the effect due to a change in the

sample. Finally, one can compute the change in the average number of hours offered by *all* women by differentiating with respect to b_0 the product of equation (10) and the probability that the average woman works.

2 In the sense that plim($W'v/n$) will typically not equal zero, where W is the matrix of explanatory variables, v is the associated error vector, and n is the number of observations.

3 Note that γ — β_1 in Heckman's (1974) notation — is missing from the cross product term in the definition of G on p. 693; this mistake is purely typographical.

4 Certainly if one wished to calculate means and variances of a variable such as income from this sample, the weights should be employed. However, when one is estimating the relationship between variables (*e.g.* estimating parameters of the Heckman model), the situation is quite different; if the weighting scheme is independent of the error structure of the model, then the weights can be ignored. If the sample can be treated as random for the purpose of estimation, then the testing of hypotheses about parameters is considerably simplified, since the properties of maximum likelihood estimates would be preserved. We know of no way to establish conclusively whether or not the error structure is independent of the weighting scheme, nor can we think of any reason why the two should be dependent. Hence we simply regarded the sample for our purposes as being truly random. It should be emphasized that we do not know of any way to estimate the variances of the parameters, and hence test hypotheses, unless the sample is assumed to be random.

It is possible to rewrite the likelihood function (11) in a manner which explicitly involves the weights; one simply regards an observation with weight w_i as being observed w_i times instead of once. The resulting function is of course not a proper likelihood function. The estimates of the parameters are readily interpretable, but their estimated variances are not. In fact, although it is not strictly proper to do so, if one formulated a null hypothesis that the true, unweighted, parameter values were equal to the values of the parameter estimates based on the weighted sample, and regarded this maintained hypothesis as being exogenous to the data (which is of course not strictly true), then the null hypothesis could not be rejected by an asymptotic likelihood ratio test. This very approximate test is certainly not a test of independence of the error structure and the weighting scheme, but does signify that the two estimates are not very different.

The elimination of observations for 847 women because of problems with the data is more troublesome. We wish that we could guarantee that their exclusion had no effect on the results, but we know of no way to reconstruct missing or contradictory data.

5 Heckman (1980) indicates that the same factors that determine work experience also determine labour force participation. He obtained significantly different coefficients in labour supply and wage functions when women's labour force experience was made endogenous by use of standard instrumental variables techniques. We have treated the actual work experience as exogenous, though a more complete model would allow this variable, as well as others such as education, to be endogenous. Usually female labour supply functions are estimated for married women only (normally with spouse present). Such restriction of this sample might induce sample selection bias. Hence, a further extension would be a model in which a nuptiality equation would be incorporated.

6 The elasticities we computed as $\frac{\partial h}{\partial W}\frac{W}{h} = \frac{W^{\lambda-1}}{\gamma h}$, , where w and h are the sample means of wages and hours of working women and λ and γ are replaced by their estimated values. These elasticities should be interpreted as structural elasticities. Other elasticities can be computed according to the discussion in Footnote 1.

7 Recall that the age-at-first-birth dummy variables measure the *incremental* effect over the constant, which captures the effect of age at first birth below 18.

8 Card and Wise (1978) have used the Project Talent data in an imaginative way to present a persuasive argument that motherhood truncates educational attainment.

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Appendix

Let $\phi(\ )$ and $\Phi(\ )$ denote the density and distribution functions of the standard univariate normal. Let Z have a standard normal distribution. Then it is easily established by integration that

$$E[Z|Z>a] = \frac{\int_a^\infty z\phi(z)dx}{1-\Phi(z)} = \frac{\phi(a)}{1-\Phi(a)} = \lambda(a) \tag{A1}$$

It follows that if X is distributed normally with mean 0 and standard deviation σ_x,

$$\begin{aligned} E[X|X>b] &= E[X|X/\sigma_x>b/\sigma_x] \\ &= \sigma_x E[X/\sigma_x|X/\sigma_x>b/\sigma_x] = \sigma_x E[Z|Z>b/\sigma_x] = \sigma_x \frac{\phi(b/\sigma_x)}{1-\Phi(b/\sigma_x)} \end{aligned} \tag{A2}$$

This result establishes that the conditional expectation of the error term in the hours equation cannot be zero unless the distribution of the error term $\mu_i - \epsilon_i = \rho_i$ is degenerate with variance zero.

It is well-known (see Mood and Graybill, 1963:202) that if X and Y are jointly normally distributed with means μ_x and μ_y, standard deviations σ_x and σ_y, and covariance σ_{xy}, that the conditional distribution of Y given $X=x$ is normal with mean

$$\mu_y + \frac{\sigma_{xy}}{\sigma_x^2}(x - \mu_x) \tag{A3}$$

In our case, μ_y and μ_x are zero, so that the mean collapses to

$$\frac{\sigma_{xy}}{\sigma_x^2}x \tag{A4}$$

We next use the fact that $E[Y] = E[E[Y|X]]$ to find

$$\begin{aligned} E[Y|X>b] &= E[E[Y|X] \text{ given } X>b] = \frac{\sigma_{xy}}{\sigma_x^2} E[X|X>b] \\ &= \frac{\sigma_{xy}}{\sigma_x}\frac{\phi(b/\sigma_x)}{1-\Phi(b/\sigma_x)} = \frac{\sigma_{xy}}{\sigma_x}\lambda(b/\sigma_x) \end{aligned} \tag{A5}$$

This result establishes that the conditional expectation of the error term in the wage equation cannot equal zero unless the covariance between μ_i and $\mu_i - \epsilon_i$ $(\sigma_{\mu\rho})$ equals zero.

It should be noted that $\lambda(a)$ cannot equal zero unless $a=-\infty$; but if $a=-\infty$, all women in the sample will work, and there can be no selection bias. Hence, in a sample of working and non-working women, a zero variance of ρ_i and a zero covariance between ρ_i and μ_i are necessary and sufficient for the conditional error terms in equations (8) and (7), respectively, to vanish.