ERODING THE ECONOMIC FOUNDATIONS OF MARRIAGE AND FERTILITY IN THE UNITED STATES*

T. Paul Schultz

Yale University

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ABSTRACT

Two explanations for the decline in the family in industrially advanced countries are increased state support of poor unmarried mothers, and increased wage earning opportunities for women relative to those available for men. This paper tests both of these hypotheses by estimating a model for the joint determination of wages of women and of their potential husbands, the probability that women are currently married, and how many children they have. The United States 1990 Census microdata sample is analyzed in combination with information on the welfare support system in the 50 different states of residence. Black and white women are considered separately, in four different age brackets. The empirical evidence is consistent with both hypotheses within both racial groups. The secular trend of increased wage rates of women relative to those of men can explain a substantial share of the observed decline in marriage and fertility, whereas the estimates suggest that variation in the generosity of welfare programs can explains relatively little of the cross sectional variation in marriage and fertility rates. In fact, the real value of these program benefits declined during the 1980s in the United States, and thus according to the cross sectional estimates reported here, these retrenchments in welfare programs should have contributed to minor increases in both marriage and fertility. More research is needed to understand the origins of the rise in women's wages relative to men's, and hence to provide a basis for forecasting future wages trends. This might provide insight into the more fundamental forces driving much of the changes in family structure noted today in high income societies. Substitution in household production as well as consumption is likely to occur in response to the widespread increase in the value of women's time relative to men's time. The increased reliance on market-purchased child care illustrates one way to reduce woman's time-intensity of both marriage and childrearing, and might reduce the negative effect on marriage and fertility of the continuing closure of the gender gap in wages.

I. Introduction

Two explanations are often offered for the decline in marriage and lifetime fertility in high-income industrialized societies. The first involves state interventions to reduce poverty among dependent children and their mothers. In this instance, the social objective is to shelter these economically vulnerable groups, but it is thought that these welfare programs have the unintended consequence of weakening the benefits from marriage and reducing the cost of childbearing among the unmarried. The second explanation involves the increase in the ratio of female to male wages for persons of comparable (marriageable) age. This trend in the relative gap in wages by gender is a consequence of many complex developments, including increases in women's education relative to men's, increases in women's persistent career attachments to the labor force, decreases in gender segregation in hiring and training by employers, and changes in the composition of production and technology that increase the derived demands for female relative to male labor. Both these general explanations for recent family change involve developments that would appear to be virtually irreversible in a modern welfare state, in the first instance, and in a society that wants to maximize its economic market return from investments in the human capital of its population, in the second instance.

There are many ways to assess the relative importance of these two developments for the prevalence of marriage and the level of fertility. As a start, the statistical association can be measured between men's and women's wage opportunities, wealth (nonearned) income, and characteristics of state welfare programs, on the one hand, and the proportion of women married and currently residing with their husband and with women's cumulative lifetime fertility, on the other hand. These cross-sectional patterns have been estimated before. The statistical methodology followed here is elaborated elsewhere (Schultz, 1994) in the context of presenting estimates based on the micro data sample of the U.S. 1980 Census of Population. Here, I extend that analysis using the 1990 Census, first to confirm the stability or change in the parameter estimates, and then to simulate how well these cross-sectional estimates explain the longer term trends over time in marriage and fertility. Of course, the goal is to gain insights into why the structure and functioning of the family has changed, and also to develop or refine forecasts about the likely evolution of the family in the future, and how society might respond if it wanted to modify that evolutionary course.

The paper is organized as follows. Section 2 outlines the conceptual framework used to explain the joint determination of wages, marriage prevalence, and fertility. Section 3 discusses the data and estimation methods. Section 4 illustrates how average conditions and behavior have changed from 1980 to 1990 in the United States among both black and white women. Sections 5, 6, and 7 report the estimates of the wage, marriage, and fertility equations, respectively. Section 8 summarizes the conclusions and identifies some directions for further inquiry.

2. The Model[1]

The decisions determining current marital status and cumulative fertility are assumed to be affected by a woman's anticipated economic opportunities, both within and outside of a marital union. Current predicted economic opportunities within and outside of marriage are used as proxies for the anticipated opportunities at the time when the decisions determining

[1] This section is parallel to Schultz 1994 section IV.

each woman's current marital status and fertility were actually made. Public welfare benefits in the United States are treated as potentially available only outside of marriage. Also, women are assumed to be able to count on income from a husband only when currently married. In other words, the likelihood that divorced women would receive child support and alimony payments from their previous husbands and that currently married women would receive welfare benefits, such as food stamps or Medicaid, are disregarded in this conceptual section for simplicity, but allowing for these possibilities would not alter the later empirical formulation of this study.

Earnings Opportunities

Both the expected hourly wage rate a woman would receive if she worked (W_f) and the expected wage rate available to her potential husband (W_m) are modeled as functions of a vector of characteristics of the woman (Y_f). Using only the woman's characteristics as explanatory variables for W_f and W_m avoids including the husband's own productive characteristics that are jointly determined together with the woman's marital status by unobservable attributes. In other words, the characteristics of observed husbands are a result of the matching process of the marriage market, and the quality of the match would tell us about (i.e. be correlated with) the unobserved qualitative characteristics of the woman that would seem likely to be relevant to her marital status and fertility decisions. Consequently, the actual husband characteristics are endogenous to my problem or correlated with the error terms in the marriage and fertility equations, and cannot be directly used here to predict potential husband wages. The statistical projection of the potential husband's wages is based, therefore, only on the woman's exogenous characteristics and those of her resident community that might affect labor market and marriage market opportunities. Thus the wage equation for the woman can be represented as

(1) $W_f = W_f(Y_f;v_f)$,

and the wage equation for the woman's potential husband is

(2) $W_m = W_m(Y_f;v_m)$,

where v_f and v_m are error terms.

The wage functions are corrected for sample selection bias (Heckman, 1987), because the woman's and potential husband's wages are observed only for some women who are wage earners and for women married to husbands who are wage earners, respectively. The estimated wage equations are used to calculate predicted values for the wage offers that would be received by each woman if she worked ($\hat{W}_f$), and by the potential husband if she were married and he worked ($\hat{W}_m$). The actual wage variables are expected to be correlated with the error terms in the marriage and fertility equations because omitted factors entering the wage function are likely to also affect marriage and child bearing decisions. It is assumed that the predicted wage variables are not correlated with the errors in the marriage and fertility equations.

The Probability of Marriage

The probability of a woman being currently married (M_f) is assumed to be a function of the predicted wage rate the woman could expect if she were employed ($\hat{W}_f$); the predicted wage rate of her potential husband ($\hat{W}_m$); a vector of welfare benefits that would be available to her if she were not married (B_f); and a vector of other variables that affect the expected resources available to the woman regardless of whether she is married or unmarried (X_f), such as her own personal wealth. The error term is denoted by u_1.

Thus, the probability of marriage is represented as

(3) $M_f = M_f(\hat{W}_f, \hat{W}_m, X_f, B_f; u_1)$.

Empirical evidence and several theories suggest that $dM_f/d\hat{W}_f < 0$, $dM_f/d\hat{W}_m > 0$, and $dM_f/dB_f < 0$ (Becker, 1974, 1981; Schultz, 1981; Wilson, 1987). These predictions are generally derived from models of marriage that hypothesize that the economic gains to marriage are enhanced by specialization in which the higher-wage individual concentrates on market work and the lower-wage individual concentrates on home production activities, including perhaps child care. I will estimate the statistical relationship represented by the marriage equation (3) using a probit model (Maddala, 1983) that assumes the error, u_1, is normally distributed and orthogonal to the observed covariates and is independently and identically distributed. The marriage probability function (3) is estimated also by ordinary least squares for a linear representation of the model, but the resulting standard errors of the coefficients are then not consistent. The standard errors for the probit specification estimated by maximum likelihood methods are consistent and should be consulted for hypothesis testing.

Fertility Equations

For both married and unmarried women, it is assumed that an increase in a woman's wage raises the shadow costs of children and increases her potential income constraint, with the hypothesized (but not theoretically necessary) net consequence of reducing her demand for births. Virtually all economic studies of fertility find that either the female wage or female education is negatively correlated with fertility. Increases in the expected wage for a potential husband are expected to exert a less negative effect on a woman's fertility than increases in her own wage (Mincer, 1963, 1985; Schultz, 1981). Indeed, under plausible assumptions regarding the father's time and market-good intensity of children it can be shown that increases in the expected husband wage should have a positive effect on fertility; there is also considerable supporting empirical evidence for this pattern (Schultz, 1981, 1986, 1994).

Therefore, for married women the fertility function is specified as

(4) $F^m = F^m(\hat{W}_f, \hat{W}_m, X_f; u_2)$,

where u_2 is an error term. It is hypothesized that $dF^m/d\hat{W}_f < 0$ and $dF^m/d\hat{W}_m > 0$. For unmarried (single) women the fertility equation is

(5) $F^s = F^s(\hat{W}_f, \hat{W}_m, X_f, B_f; u_3)$,

where u_3 is an error term and where the vector of welfare variables (B_f) is introduced. It is hypothesized that $dF^s/d\hat{W}_f < dF^m/d\hat{W}_f < 0$, $dF^m/d\hat{W}_m > dF^s/d\hat{W}_m > 0$, and $dF^s/dB_f > 0$ (Becker, 1981; Leibowitz et al., 1986; Rosenzweig and Schultz, 1985; Schultz, 1981, 1986, 1990; Whittington and Peters, 1990). Increasing the market wage opportunities for woman is thus hypothesized to reduce her demand for children by a larger amount if she is unmarried than if she is married, presumably because having a husband to coordinate the allocation of market and nonmarket activities allows the married woman to sacrifice relatively less resources to have an additional child. Correspondingly, the positive impact of market wage opportunities for her potential husband on her demand for children will be larger if she is actually married than if she is currently unmarried and in a weaker position to enforce her claims for child support on the father. These are plausible, but also strong assumptions, needed to obtain clear predictions on how female and male wage opportunities will influence a woman's expected fertility, independent of her current marital status.

Since marital status can be endogenously changed by a woman, an equation is needed for a woman's expected fertility, (F^t), without conditioning on her marital status. Unconditional fertility is the sum of the marital-status-specific fertilities weighted by the respective probabilities of being currently married or unmarried; that is,

(6) $F^t = M_f F^m + (1 - M_f)F^s.$

If equation (4) and (5) were estimated to assess directly the factors affecting fertility of married and unmarried women separately, sample selection bias could be expected, because the error, u_1, in the marriage selection equation (3) is likely to be related to errors, u_2 and u_3, in the marital-status-specific fertility equations. Common omitted variables are likely to affect all three outcomes. I see no cogent conceptual basis for identifying the required sample selection correction procedure (Heckman, 1987), and therefore I estimate the linearized reduced form equation (6) for the combined married and unmarried samples. This estimation approach is designed to eliminate bias from sample selection or the endogeneity of marriage and the matching it facilitates of men and women on unobserved characteristics. Fortunately, some of the signs of the derivatives of the unconditional reduced-form fertility function (6), ignoring potential higher order interaction terms, are implied by the conditions already hypothesized for equations (3), (4) and (5), and these can be tested empirically.

The effect of the woman's predicted wage on unconditional fertility are expected to be negative, and the effect of the predicted wage of her potential husband to be positive, provided that the expected married fertility level is higher than the unmarried fertility level ($F^m - F^s > 0$). That is,

(7) $dF^t/d\hat{W}_f = (dM_f/d\hat{W}_f)(F^m - F^s)+(dF^m/d\hat{W}_f - dF^s/d\hat{W}_f)M_f+(dF^s/d\hat{W}_f) < 0,$

and

(8) $dF^t/d\hat{W}_m = (dM_f/d\hat{W}_m)(F^m - F^s)+(dF^m/d\hat{W}_m - dF^s/d\hat{W}_m)M_f+(dF^s/d\hat{W}_m) > 0.$

Welfare benefits are expected to reduce the probability of a woman being married, because poor families cannot normally qualify in the United States for welfare benefits when an able-bodied husband is present. Welfare benefits may thus indirectly tend to reduce fertility, provided that the expected married fertility level exceeds the expected unmarried fertility level, as assumed earlier. At the same time, however, by relaxing the budget constraint for unmarried mothers, welfare benefits would tend to encourage childbearing among unmarried women. Thus, a priori information from this model of marriage and fertility is insufficient to prescribe the sign for the total effect of welfare benefits on the unconditional fertility level of a woman. That is,

(9) $dF^t/dB_f = dM_f/dB_f(F^m - F^s) + dF^s/dB_f \gtrless 0.$

The second term on the right side of equation (9) is the child price subsidy effect of welfare benefits that is often emphasized as possibly encouraging unmarried women to have additional children (Murray, 1984), whereas the first term is negative to the extent that expected fertility of married women exceeds that of unmarried women, which it does in the United States. Studies of the effect of welfare systems on fertility often analyze only the impact of welfare regime on the fertility of unmarried women, estimating only the second term's positive effect. Including also welfare's potential effect operating through changes in marital status is likely to reduce the estimated unconditional effect of welfare on fertility, and perhaps even change its sign.

Consequently, only an empirical analysis of the unconditional effects of welfare benefits on average fertility of all women can resolve whether a particular scheme of state subsidies for the children of unmarried women are increasing or deceasing the overall level of childbearing in a society. In the U.S. context, the provision of medical insurance for a poor mother and her child is also an increasingly important part of the public welfare program that has in the past benefited primarily unmarried mothers. Many states in legislating their package of health care for welfare-assisted women, or Medicaid, often include subsidized access to birth control services and supplies that would be expected to also lower fertility.

3. Data and Estimation Methods

One-in-every-hundred non black (white) woman age 15-65 is sampled from the one-in-a-hundred public use microdata sample of the 1990 Census of Population.[2] One-in-ten black woman of the same age bracket is sampled to obtain black and white samples of approximately the same size. Information was retained on a woman's husband if he was also enumerated by the Census as residing in her household, and his co-residence defines whether she is subsequently treated as "currently married". Table 1 reports the means and standard

[2]The random sample was extracted for me by CIESIN with the assistance of Paul McGuire.

TABLE 1

MEANS AND STANDARD DEVIATIONS[a] OF THE VARIABLES
FOR BLACK AND WHITE WOMEN IN 1990 CENSUS, AGE 15-24, 25-34, 35-44, 15-65 [a]

	Black				White			
Variables	15-24	25-34	35-44	15-65	15-24	25-34	35-44	15-65
Dependent Variables:								
Currently Married and Husband Present (=1)	.110	.330	.431	.355	.342	.664	.742	.672
Children ever born	.872 (.0104)	1.71 (1.42)	2.33 (1.72)	2.32 (2.76)	.574 (.889)	1.34 (1.22)	1.95 (1.39)	1.96 (1.70)
Explanatory Variables:								
Woman's Predicted Log Hourly Wage	1.72 (.215)	1.71 (.431)	2.26 (.320)	1.72 (.524)	1.50 (.426)	2.04 (.249)	2.00 (.375)	1.94 (.347)
Potential Husband's Predicted Log Hourly Wage	1.69 (.373)	1.63 (.487)	1.85 (.506)	1.74 (.476)	1.90 (.404)	2.14 (.567)	2.43 (.593)	2.14 (.541)
Woman's Property Income ($1,000 per year)	.0532 (1.04)	.0336 (.616)	.0495 (.394)	.0748 (.831)	.160 (2.17)	.175 (1.71)	.474 (3.31)	.609 (3.65)
Schooling (years completed)	11.7 (1.95)	12.3 (2.39)	12.4 (2.58)	11.9 (2.75)	12.0 (2.31)	13.1 (2.54)	13.1 (2.87)	12.6 (2.82)
Sample size	924	1973	1731	6600	884	2210	2127	8123

[a] Standard deviations are shown in parentheses. The predicted wage variables are obtained from sample selection corrected wage equations as reported in Tables A-1.

deviations of the central variables in the study for black and white women age 15-24, 25-34, 35-44, and 15-65.

The empirical analysis proceeds in several stages. For each race and age group of women, a probit equation is estimated by joint maximum likelihood methods to explain the probability that the woman is a wage earner and to predict her hourly wage, measured in logarithms (Maddala, 1983; Heckman, 1987). These estimates are reported in Table A-1. A parallel wage function is estimated for husbands jointly with the probit equation for the probability that the woman is both married and her husband reports a wage (Table A-1). As noted earlier, the characteristics of the woman are used to predict the probit and wage functions for her husband, thereby avoiding match-specific husband variables that are jointly determined aspects of the marital union and hence are likely to be correlated with the error in the wage, marriage, and fertility functions. This round-about procedure, as noted earlier, is designed to avoid simultaneous equation bias.[3]

The predicted wage variables for each woman and her potential husband are then included first in the probit equations for the probability that a woman is currently married (and co-residing with her husband)(Table 3), and second in an ordinary least square equation for the number of children ever born (Table 4). The linear probability model for the marriage

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[3]For example, a woman who had an unexplained demand for a large number of children might be expected to have a tendency to lower her (wage) standards for a husband to obtain her fertility goal. A lower actual husband wage than that predicted on the basis of her observed characteristics would then be associated with a higher than explained fertility level. If we simply used the husbands' characteristics to predict their wages, the tendency would be to attribute, in this example, a smaller positive potential husband wage effect on her fertility, because the husband characteristic-based prediction of his wages would also include an negative effect of the woman's preferences for children on the market wage quality of her marital match.

TABLE 2

LINEAR PROBABILITY MODEL ESTIMATES OF LIVING WITH A HUSBAND, FOR WOMEN, BY RACE AND AGE[a]

	Black				White			
Explanatory Variables	15-24	25-34	35-44	15-65	15-24	25-34	35-44	15-65
Woman's Predicted Log Hourly Wage [b]	-.208 (2.59)	-.0322 (1.69)	-.126 (4.89)	-.122 (12.5)	-.247 (6.48)	-.268 (7.83)	-.219 (10.8)	-.300 (21.8)
Potential Husband's Predicted Log Hourly Wage [b]	.298 (9.26)	.665 (40.1)	.764 (46.1)	.697 (67.9)	.442 (12.2)	.498 (33.9)	.511 (33.1)	.499 (55.3)
Woman's Property Income per year (x10 $^{-3}$)	-.118 (1.04)	-2.13 (1.70)	-38.1 (1.87)	-6.56 (1.25)	-1.98 (.30)	-6.67 (1.40)	-.615 (2.64)	-.695 (5.87)
Age (years)	-.104 (1.15)	.0041 (.07)	-.371 (4.25)	.0082 (3.20)	.25 (2.05)	.0949 (1.44)	-.0554 (.66)	.0228 (8.62)
Age Squared (x10 $^{-2}$)	.374 (1.70)	.0160 (.15)	.445 (4.02)	-.00272 (.88)	-.453 (1.55)	-.167 (1.51)	.0078 (.73)	-.0197 (6.26)
Hispanic (=1)	.0296 (.30)	.112 (2.41)	.131 (2.16)	.064 (2.07)	.0207 (.46)	-.0176 (.56)	.0694 (2.07)	-.0023 (.13)
AFDC Benefits Indicator ($ per month x10 $^{-2}$)	.00277 (.25)	-.0352 (4.85)	-.0296 (3.97)	-.0282 (6.81)	-.0458 (3.52)	-.0438 (6.26)	-.0196 (2.91)	-.0258 (6.87)
Medicaid Expenditures ($ per family per month x10 $^{-2}$)	-.0558 (2.14)	-.0131 (2.60)	-.0068 (.37)	-.0271 (2.65)	-.0119 (.45)	-.0495 (3.16)	-.0111 (.75)	-.0026 (.32)
Intercept	.619 (.66)	.791 (.86)	7.02 (4.11)	-.797 (17.0)	-3.25 (2.56)	-1.13 (1.17)	1.04 (.64)	-.275 (5.62)
R^2	.140	.478	.558	.451	.217	.357	.345	.329
Sample size	924	1963	1731	6600	884	2210	2127	8123

[a] Absolute value of asymptotic t ratio is reported in parentheses probit coefficient.
[b] Predicted log wage variable on the basis of the appropriate selection-corrected wage equation in Table A-1.

TABLE 3

PROBIT ESTIMATES OF LIVING WITH A HUSBAND, FOR WOMEN, BY RACE AND AGE[a]

Explanatory Variables	Black 15-24	Black 25-34	Black 35-44	Black 15-65	White 15-24	White 25-34	White 35-44	White 15-65
Woman's Predicted Log Hourly Wage [b]	-1.32 (2.58)	-.0286 (.32)	-.229 (1.82)	-.412 (9.54)	-.806 (6.28)	-.831 (6.14)	-.886 (9.25)	-1.13 (19.5)
Potential Husband's Predicted Log Hourly Wage [b]	2.22 (11.3)	2.11 (27.5)	2.54 (27.9)	2.34 (47.0)	1.44 (11.2)	1.52 (26.4)	1.91 (24.5)	1.77 (44.7)
Woman's Property Income per year ($\times 10^{-3}$)	-57.7 (.51)	-157. (1.40)	-110. (1.15)	-15.6 (.73)	-67.2 (.25)	-23.5 (1.52)	-20.0 (2.20)	-23.4 (5.55)
Age (years)	.364 (.47)	.166 (.55)	-1.58 (3.68)	.087 (7.26)	1.20 (2.42)	.367 (1.41)	.327 (.84)	.0983 (9.87)
Age Squared ($\times 10^{-2}$)	-.089 (.05)	-.170 (.33)	1.93 (3.54)	-.068 (4.81)	-2.36 (2.02)	-.619 (1.41)	-.457 (.92)	-.0884 (7.49)
Hispanic (=1)	-.749 (1.25)	.428 (2.02)	.471 (1.80)	.282 (2.21)	.056 (.37)	.117 (.99)	.261 (1.80)	-.0291 (.45)
AFDC Benefits Indicator ($ per month $\times 10^{-2}$)	-.0176 (.29)	-.126 (3.64)	-.119 (3.14)	-.104 (5.63)	-.143 (3.27)	-.158 (5.67)	-.0928 (3.00)	-.0970 (6.54)
Medicaid Expenditures ($ per family per month $\times 10^{-2}$)	-.394 (2.28)	-.218 (2.39)	.0480 (.53)	-.126 (2.73)	-.0408 (.46)	-.188 (2.99)	-.0310 (.46)	-.0028 (.09)
Intercept	-9.88 (1.21)	-6.74 (1.51)	28.2 (3.34)	-5.63 (22.9)	-16.4 (3.13)	-6.17 (1.62)	4.42 (.58)	-3.03 (16.0)
Chi Squared Statistic (8 df.)	217.0	1004.	1113.	3408.	214.8	839.1	802.2	3078.

[a] Absolute value of asymptotic t ratio is reported in parentheses probit coefficient.
[b] Predicted log wage variable on the basis of the appropriate selection corrected wage equation in Table A-1.

TABLE 4

ORDINARY LEAST SQUARES REGRESSIONS FOR CHILDREN EVER BORN TO WOMEN, BY RACE AND AGE OF THE MOTHER[a]

	Black				White			
Explanatory Variables	15-24	25-34	35-44	15-65	15-24	25-34	35-44	15-65
Woman's Predicted Log Hourly Wage [b]	.288 (1.29)	-.872 (11.6)	-.361 (2.73)	-.853 (16.5)	-.854 (11.8)	-1.84 (18.6)	-.990 (13.1)	-1.28 (24.0)
Potential Husband's Predicted Log Hourly Wage [b]	-.678 (7.57)	-.0602 (.92)	-.140 (1.65)	-.103 (1.90)	.0225 (.33)	.292 (6.86)	.185 (3.20)	.207 (6.13)
Woman's Property Income per year (x10 $^{-3}$)	-.0356 (1.13)	-.0514 (1.04)	-.114 (1.09)	-.0882 (3.17)	-.0059 (.47)	-.0210 (1.53)	-.0101 (1.16)	-.0132 (2.99)
Age (years)	.488 (1.94)	.126 (.51)	-.472 (1.06)	.212 (15.8)	.655 (2.82)	.172 (.90)	.365 (1.17)	.204 (20.6)
Age Squared (x10 $^{-2}$)	-1.00 (1.64)	.0411 (.10)	.711 (1.25)	-.160 (9.78)	1.17 (2.11)	-.113 (.35)	-.407 (1.02)	-.164 (13.9)
Hispanic (=1)	-.112 (.42)	-.045 (.25)	.755 (2.34)	.0879 (.54)	.197 (2.31)	.201 (2.21)	.644 (5.14)	.437 (6.75)
AFDC Benefits Indicator ($ per month x10 $^{-2}$)	-.0501 (1.62)	-.0855 (2.99)	-.122 (3.19)	-.105 (4.83)	-.0449 (1.82)	-.0181 (.89)	.0170 (.68)	.00754 (.54)
Medicaid Expenditures ($ per family per month x10 $^{-2}$)	-.102 (1.30)	-.0991 (1.33)	-.0969 (1.04)	-.158 (2.92)	-.0553 (1.10)	.0318 (.70)	.0430 (.78)	.0345 (1.12)
Intercept	-3.97 (1.53)	.323 (.09)	11.4 (1.30)	-1.04 (4.21)	-6.60 (2.74)	.368 (.13)	-4.68 (.76)	-1.39 (7.61)
R^2	.103	.108	.0394	.250	.200	.187	.098	.284

[a] Absolute value of asymptotic t ratio is reported in parentheses probit coefficient.
[b] Predicted log wage variable on the basis of the appropriate selection corrected wage equation in Table A-1 for age 15-65 and analogous estimates for other age groups.

equation is also reported in Table 2 for comparison, although as noted caution should be exercised in interpreting tests of statistical significance. The common practice of examining fertility within marital status groups is not justified if the error in the marital status equation, u_1, is correlated, as seems plausible, with the error in the marital status-specific fertility equations, u_2 and u_3. The fertility equation is therefore estimated unconditionally on marital status for the reasons outlines in the prior section.

4. Change in Family Conditions

In the decade of the 1980s, black women age 15-65 experienced an increase in their market-offered wage rates of 48 percent, according to the estimates of sample-selection-corrected wage functions (Tables 1 and A-1; Schultz, 1994, Tables 2 and A-1). White women age 15 to 65 experienced an even faster growth in wages of 54 percent from 1979 to 1989. The wages of the prospective husbands of these women grew more slowly in this decade, by 27 percent for white women and 23 percent for black women. Thus, a black women's wage relative to that of her prospective husband, increased from 73 to 98 percent, reaching virtual gender parity by 1989, according to these estimates.[4] The change in white women's wages relative to those of their prospective husband's was also dramatic, increasing from 53 to 80 percent of gender parity. Of course, women continue to work somewhat fewer hours than husbands, but they appear to be catching up to male levels in this area as well.

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[4]More precisely, this figure is obtained by subtracting from one the difference in the mean logarithmic wage rate for potential husbands and women (Table 1).

Property income--dividends, interest, royalties, and rents--of women increased more rapidly than their wage rates but nonetheless constituted a small fraction of their income, particularly for blacks. The mean annual property income of a black woman age 15-65 doubled from 1979 to 1989 according to the U.S. Census, but was still only $75 in the latter year (Table 1; Schultz 1994, Table 2). For comparably aged white women, property income slightly more than doubled in this decade, but represented on average $609 by 1989. Both the rise in women's wages relative to men's and the increase in property income of women are expected to diminish the economic motivation for marriage among women, and at least the increase in women's wages is expected to reduce fertility.

Public transfers programs in the United States have also changed over time the levels of support provided poor mothers and their dependent children. Benefits under the U.S. Welfare Program called Aid to Families with Dependent Children (AFDC) vary by state, and coordinate payments of cash and food stamps. The maximum benefit in cash and food stamps for a lone mother with one child is considered here to be one comprehensive indicator of the generosity of a state's welfare system. This state benefit was $265 per month for the average black women age 15-65 in my sample in 1989 and $311 for an average white women, given the different racial distribution of the population across states. The Medicaid reimbursements in 1989 for medical services provided per AFDC adult recipient and per AFDC child recipient is divided by the proportion of AFDC families in each state receiving Medicaid. This is the second indicator I shall use for the generosity of the state welfare system. This measure of the expected Medicaid reimbursements of doctor and hospital expenses per AFDC family of two is equal to $162 per month for black women in 1989 and $170 for white women in my sample, where again the race difference is due to the different geographical distributions of the groups across the states. This measure of AFDC cash and food benefits increased nominally about 15 percent from 1979 to 1989, whereas the available measures of Medicaid reimbursements increased in nominal terms 29 percent. The U.S. consumer price index increased in this time period by one-half, and the cost of medical care increased even more rapidly. Thus, the real value of these welfare programs for poor mothers in the United States declined markedly in this decade while the share of all benefits from Medicaid increased (Moffitt, 1992).

In the 1980s the proportion of black women currently married and residing with their husband, according to the census enumeration, decreased within the three childbearing age groups, but declined mostly between the ages 25 and 44 (Table 1; Schultz, 1994, Table 2). Among whites the proportion currently married increased in the 1990 census sample for women age 15-24, but declined somewhat for comparisons at later ages. The major contrast remains the difference in proportion currently married between the races, where black marriage prevalence rates are about half the level reported for whites (Bennett et al., 1989).

The number of children ever born has increased for women age 15-24 from .56 to .87 for blacks from 1980 to 1990, and from .13 to .57 for whites. The number of children ever born decreased slightly at ages 25-34 from 1.87 to 1.71 for blacks, and from 1.38 to 1.34 for whites. Among women age 35-44 the more marked declines in fertility that started during the 1960s are evident, with fertility falling from 3.21 to 2.33 over the decade for blacks, and from 2.58 to 1.95 for whites. Although period-specific birth rates have been relatively stable in the 1970s and 1980s, these cohort fertility rates suggest an increase in early childbearing has taken place. Although not the focus of this study, it should be recalled that 20 percent of white births by 1990 are occurring out-of-wedlock, whereas the proportion among black births is 60 percent by that date (Schultz, 1994).

5. Wage Estimates

The wage function estimates for black and white women age 15 to 65 are reported in Appendix Table A-1, and analogous estimates are calculated in the three age subsamples but are not reported to save space. Several features of these estimates may be noted.

The sample selection correction procedure is identified by several variables that are theoretically expected to reduce the individual's propensity to participate in wage employment and perhaps also reduce her likelihood of being married, but are not expected to affect directly market wage opportunities of the individual. As seen in Table A-1, these identifying exclusion variables that enter only the wage-earner probit equation are the woman's property income, the squared value of this property income to reflect nonlinear effects, and three state-level welfare program measures: AFDC benefits, Medicaid reimbursements, and duration of unemployment benefits.[5] The identifying variables are jointly significant according to joint likelihood ratio tests at the .05 percent level for white women, but only at the 12 percent level for black women.[6] The estimated correlation of the errors in the wage-earner probit and wage

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[5]From Table A-1 it may be observed that property income is a significant factor in reducing participation as a wage earner for white women and their prospective mates. Duration of unemployment benefits reduces wage participation of black women and their prospective husbands and white women as well. AFDC benefits lower wage participation rates in all four groups. Medicaid benefit levels reduce the wage participation of black women's prospective husbands. In the marriage and fertility equations these wage functions are reestimated within the ten-year age subsamples with few substantive changes.

[6]The likelihood ratio is distributed as a chi-squared statistic with 5 degrees of freedom for women in Table A-1. It is 8.75 for black women, which is significant at the 12 percent critical level. For

functions, or rho, is statistically significant in all four estimations, indicating that sample selection bias may be present (Heckman, 1987). Unobservables that contribute to an individual's probability of being a wage earner (being married) are positively associated with that individual receiving a higher market wage.

The coefficients in the wage function on years of schooling indicate that from 1979 (Schultz, 1994) to 1989, these estimates of the private returns to secondary and higher education increased, and they are higher for black women than for white women in 1989. This change in U.S. wage structure during the 1980s has been widely noted for males (e.g., Katz and Murphy, 1992) and documented also for females (Bound and Johnson, 1992). An extra year of secondary schooling is associated in 1989 with black women receiving 17 percent higher wages, whereas a year of higher education earned a black woman a 16 percent higher wage. Ten years earlier, the returns were lower, 11 percent at both levels of education (Schultz, 1994). The wage premium in 1989 for a year of secondary and higher education is 10 and 13 percent for white women, compared with 6 and 11 percent in 1979. In none of the wage functions is there a significant return to primary schooling, but relatively few younger individuals end their schooling before the eight grade, and therefore it is not surprising that these returns to primary schooling are imprecisely estimated in this population.

The potential husband of a black or white women receives a higher wage if she is better educated. But the proportionate increase in her partners wage is smaller than that in her own wage. Thus, the ratio of the woman's wage to that of her potential husband increases

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white women it is 25.0, which is significant at the .02 percent level. For black and white husbands the significant levels are .02 and 5 percent, respectively.

with her education toward gender parity. If this ratio of the wage a woman can expect to earn relative to that her potential husband can expect to earn approximates the economic gains she may obtain from marriage, these wage function estimates suggest that the economic gains to marriage decline for women as their education increases. This could occur for at least two reasons. As her education increases relative to that of the pool of marriageable males, she is more likely to settle for a husband whose education is relatively lower then her own. Also as the ratio of a woman's wage to her potential husband's wage approaches parity, the economic gains from trade within marriage are eroded that arise from traditional production specialization, in which the husband works predominantly in the labor force and the wife works more in home production activities, including perhaps child care (Becker, 1981).

6. Marriage Estimates

Increasing the women's wage opportunities decreases marriage among white women in all age groups, whereas increasing the prospective husband's wage opportunities is associated with increases in the probability of being currently married. If the logarithm of the gender ratio of wages (W_f/W_m), which is equivalent to the difference in gender-specific log wage variables, were the critical factor decreasing the probability of being married, then equal but opposite-signed coefficients would be expected on the female and male log wage variables in the marriage equation. This possible additional restriction to the model estimated here is soundly rejected for both blacks and whites in either the linear probability model (Table 2) or the probit model (Table 3). For white women the absolute magnitude of the male wage coefficient is about twice as large as the (negative) female wage coefficient in the marriage equation. Among black women the effect of male wage opportunities is six or more times larger than that of female wages, except for the youngest age group where they are more similar. A ten percent increase in black male wages, holding constant black female wages, is associated with an increase in marriages for black women age 15 to 65 of seven percentage points, or by one-fifth (.07/.36). If this ten percent gain in black male wages occurred in conjunction with a comparable ten percent increase in female wages, the simulated effect would still be a 5 percentage point rise in marriage prevalence. For white women an equiproportional increase in wages for women and their prospective husbands would also increase the proportion currently married, but by only about 2 percentage points, or by one-thirtieth (.02/.67). Evidently, black marriage rates today are more sensitive than white marriage rates to the gender wage gap and are particularly responsive to changes in the wage opportunities available to men who are prospective husbands.

Another issue to explore with these estimates is how much the prevalence of marriage would be expected to change if women's wages caught up to those of their prospective husbands, or the gender wage gap closed. For black women age 15 to 24 and 25 to 34 the wage gap is estimated to be only 3 and 8 percent, whereas for whites in these age groups the gender wage gap still remains 40 and 10 percent, respectively (Cf. Table 1). If women's wages were to rise to the prospective husband's level, black marriage rates from age 15 to 24 and 25 to 34 would fall by 0.6 and 0.3 percentage points, respectively. For whites the implied larger increase in the woman's wage opportunities would be associated with a relatively large 9.9 percentage point decline in marriage at ages 15 to 24. Currently, 34.2 percent of this age-race group is married. At age 25-34 the closure of the white gender wage gap is associated with a smaller 2.7 percentage point decline in marriage from the current level of 66.4 percent. Thus, the estimated effect of closure in the wage gap for whites is mostly to delay early marriage, but not necessarily reduce substantially the prevalence of marriage among older white women.[7]

An increase in AFDC benefits of $115 per month for a mother with one child, which is equivalent to a standard deviation change in this sample, is associated in the linear probability model with a 3 percentage point decline in black marriages between ages 15 and 65, from 36 to 33 percent. For white women marriage would decline by the same amount, 3 percentage points, but this represents only half the proportionate effect for whites as for blacks, because of their greater current marriage rates. The proportionate effects are larger for younger white women age 15 to 34 and for black women age 25 to 44.

An increase in Medicaid reimbursements per AFDC family of two of $53 per month, which is again a standard deviation in the sample, is associated with a decline of only 1.4 percentage points in the proportion of black women age 15-65 currently married. The effect is, however, larger for younger black women age 15-24. Among white women only those age 25-34 show a significant marriage response to Medicaid, and a standard deviation increase in Medicaid has the effect of reducing marriage prevalence in the linear probability model by 2.6 percentage points for this age group. In sum, the welfare system appears to have relatively small effects on the probability of being currently marriage compared to the effects of changes

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[7]One can also estimate the same model for the probability of ever being married to assess whether the current marriage response is due to avoiding marriage or increasing divorce. For black the wage effects are operating predominantly through avoiding marriage whereas for whites divorce or separation is residually responsive to the wage gap.

in gender-specific wage rates, but they are nonetheless statistically significant and consistent with other studies (Moffitt, 1990; Yelowitz, 1994).

7. Fertility Estimates

The labor market effects on fertility are similar for white women in 1990 as they were in 1980. The negative impact of women's wage opportunities decreased slightly in magnitude in all age groups from 1980 to 1990, but remain large and statistically significant. The prospective husband's wage opportunities tend to be a positive factor on the fertility of white women, as they were on marriage, but they are insignificant in 1990 for the youngest group of white women, age 15-24. The negative female wage coefficients remain absolutely larger than the positive male wage coefficients, implying that an equiproportional increase in the wages of men and women is still in 1990 associated with lower fertility among white women.

Among black women the effect of female and male wage opportunities appear to have changed from 1980 to 1990, at least among the younger women age 15 to 24. In 1980 the fertility of black women was lower when women's wage opportunities are more attractive, but in 1990 this pattern in fertility predicted by economic models of fertility is evident only among black women over age 25. Conversely, the wage opportunities of her prospective husband that were positively associated with fertility among black women age 15 and 34 in 1980 are in 1990 negatively associated with her fertility and these unanticipated effects are significant in the youngest age group. This instability of fertility coefficients on black men's and women's wages may be a statistical artifact of the multicollinearity in the gender specific-predicted wage rates for blacks. It should be recalled that male and female wage rates nearly converged for blacks from 1980 to 1990 and may be difficult to precisely separate in 1990.

Property income has a strong lifecycle component and may be accumulated from past wage earnings. The negative fertility coefficient on women's property income for the age aggregated samples of white and black women may thus not reflect a causal effect, if property income were viewed as endogenous. Within all age disaggregated samples the association of fertility with property income is not significant but remains negative. Property income is, however, consistently associated with lower marriage rates, although these coefficients are significant only for white women over age 35 (Cf. Schultz, 1990).

Most Hispanic women are in the white (i.e., nonblack) sample. The coefficient on the white Hispanic dummy for women in 1980 and 1990 suggests Hispanic women have about the same probability of being currently married as do other white women, given the other control variables in the model (Table 2 or 3). Hispanic white women born after 1955 (age 34 or younger in 1990) have about 0.2 more children than otherwise comparable white women. This contrasts with a larger 0.5 to 0.7 child differential between Hispanic and other white women born before 1955. The Hispanic fertility differential thus appears to be decreasing in the United States as these cohorts have become more assimilated, and this fertility pattern across birth cohorts is consistently portrayed in both the 1980 and 1990 census data (Schultz, 1994).

These empirical findings assign importance to the changing structure of wages in the U.S. labor market and specifically to the differences between the wage opportunities of men and women as a forcing factor explaining cross-sectional differences in marriage arrangements and fertility. These cross-sectional estimates in combination with changes over time in the conditioning wage variables help to account for the decline in the prevalence of marriage and fertility that has been observed across cohorts born after 1935. Although the closure in the gender gap in wages is widely noted in the world (Schultz, 1993), the factors that underlie this convergence differ in the United States, Europe, and other countries (Blau and Kahn, 1997).

In many parts of the world, women's schooling has increased much more rapidly than men's in the last several decades, and this secular increase in the educational attainment of women relative to men can be attributed a major role in closing the gender wage gap (Schultz, 1993). This is undoubtedly the case in most high income countries (Mincer, 1985). But in the United States, where women received roughly the same number of years of schooling as did men at the outset of the Twentieth Century, other explanations for the closure in the U.S. gender wage gap must be found (Goldin, 1990).

A major factor for the convergence in gender specific wages in the United States after 1980 has been the increased labor force experience of the average working women (Smith and Ward, 1985). However, the increase in women's labor force participation and hours of market work may also be caused by the fertility decline that I seek to explain here. For with fewer births, women confront lower costs of entry into the labor force and have fewer difficulties maintaining a long-term career commitment to a job. The resulting strengthened attachments to the labor force generally facilitate women's access to on-the-job training that is subsequently reflected in their increased wage rates over the lifecycle. Therefore, to avoid simultaneous equation bias, the wage functions for women were <u>not</u> predicted on the basis of their <u>actual</u> labor force experience or indirectly on their fertility and marital status, but only on their exogenous potential post-schooling experience.[8] The estimated effects of wage opportunities are thus not biased by common unobservables affecting both fertility and female labor force behavior, and consequently impacting on their current wage rates.

Alternative empirical and statistical specifications of the marriage and fertility models were also estimated, without changing notably the empirical findings emphasized in this paper. The effects of proportional (logarithmic) changes in wages or absolute property income changes may not be linear on the probability of being married or on the number of children born. Higher order polynomials (e.g. squared values) were added in these income variables throughout the analysis with little effect on the implied derivatives of the functions evaluated at sample means. These higher order terms are probably useful if estimation of effects is sought among segments of the population far from the sample mean, but multicollinearity then also becomes a greater problem as higher order terms are included as regressors. Regression models that explicitly recognize the censored restrictions on the dependent variable for fertility, i.e. fertility may not take on a negative value, leads to alternative nonlinear models, such as proposed by Tobin (1958) to study the demand for lumpy durable goods, but did not change the main results reported here.[9] It does not appear that the

[8]The experience variable included in the wage equations is simply the one proposed by Mincer (1974) of age - years of schooling completed - seven (i.e., age of entry into school).

[9]See Schultz 1994 for Tobit estimates for comparison with OLS in 1980 census study (Table A-4). This form of Tobit and ordered probit models were also estimated for the fertility model, to allow for the censored values of fertility at zero and the possibly nonlinear scaling of fertility that might be better fit across parities by using an alternative model (Maddala, 1983). Although a minor improvement in statistical fit was achieved by the Tobit for the youngest sample of women, age 15 to 24, for whom the fraction censored at zero children was largest, the derivatives of the expected value locus with respect to the conditioning variables were very similar between the Tobit and the OLS model, and parallel in statistical significance.

relationships between the female and potential husband wage opportunities and welfare variables and the marital status and fertility variables depend sensitively on which of these alternative model specifications is adopted.

8. Conclusions

This study of data from the U.S. 1990 Census reconfirms many of the pattern of determinants of marriage and fertility that has been observed in repeated U.S. cross sections (e.g., Schultz, 1981, 1986, 1994). Fewer women are currently married than in the past, a larger share of children are being born out of wedlock, and a growing fraction of children grow up without their father residing in their household. These trends are of course not unique to the United States. Their determinants in the United States are probably also operating to create similar family patterns in other high-income countries.

These changes in marriage and fertility, I have suggested, arise in part from the tendency for women's wages to increase recently more rapidly than men's wages. Many conceptual approaches to the family recognize that the lifetime nuclear family is weakened by the closure in the gender gap in wages and the loss of male employment prospects relative to those available to females in some demographic groups (e.g., Wilson, 1987). The economic motivation for women to marry and to stay married has diminished, at least insofar as the benefits of marriage are enhanced by specialization between husband and wife in market and nonmarket production activities. The secular increase in women's wages relative to men has its origins in many complex and poorly understood factors, but these factors are not likely to be readily reversed in most market-oriented economies.[10] As long as children remain privately a more costly burden on their mother's time and earnings opportunities than on their father's time, fertility is also likely to decrease in societies as the gender gap in wages narrows, other things being equal. That tendency is clearly evident in cross sectional differences analyzed in this study from the 1990 U.S. Census.

Marriage and the fertility that marriage fosters by socializing and supporting children involve adaptable social institutions that can change in response to new external conditions. Different rules may evolve for sharing the more limited economic gains from long-lasting marriages. The economic and psychological costs of childbearing and childrearing can also be partially redistributed among mothers, fathers, the local community, and the nation state, if common grounds can be found for exchange with an improved outcome for all parties. Modification of customary and individualized arrangements surrounding family functions and intrafamily allocations might halt the decline in marriage and even increase fertility within marriage and even increase fertility outside of wedlock. Alva Myrdal (1945) addressed these challenges facing the egalitarian welfare state at a time when Sweden's fertility first dipped below replacement levels in the 1930's. Have we come much further in the last fifty years in understanding the nature of this dilemma and formulating improved social policies that advance and harmonize the well-being of children, women, and men?

According to the estimates reported in this paper, black women in the United States have increased their wage rate by one-third during the 1980's, relative to that which their

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[10]The patterns of change in gender differences in wages may be different in some transition economies where female participation was nearly universal before the end of central planning.

potential husband's can earn. The gender wage gap among black Americans has virtually closed by 1990. The relative advancement of white women is no less marked, increasing their wage relative to their prospective husbands by one-half, reaching four-fifths of their potential husband's wage by 1990 (Table 1). In the cross section of the population examined here, such a sharp narrowing of the gender wage gap is associated with a substantial reduction in the proportion of women who are currently living with a husband and is linked to a decline in their level of fertility. But the downward adjustment in marriage appears to have been accentuated because the reduction in the gender wage gap came about largely due to the decline in male wage opportunities, and not a rise in female wages, particularly among black Americans (Compare Table 1 and Schultz, 1994). This is reflected in the asymmetric tendency for the same percentage (logarithmic) decline in male wage opportunities to exert a larger effect in 1990 reducing the prevalence of marriage among both white and black women, than does the same percentage increase in female wage opportunities (Table 2 and 3). With regard to fertility, with the exception of black women age 15-24, a proportionate increase in female wages continues in 1990 to be associated with a larger reduction in fertility than is a decline of the same proportion in potential husband wages (Table 4). The prevalence of marriage is particularly sensitive to the evolution of labor market opportunities for men, whereas fertility remains more sensitive to the market value of women's time, which it is widely postulated to have the major effect on the opportunity cost to women of childbearing.

Although the generosity of the state welfare system is associated with fewer women living with a husband, as expected, this statistically significant relationship can explain only a modest fraction of the decline in marriage in the last fifty years. Moreover, welfare system benefits for unmarried women with children appear to depress average fertility levels for all (married and unmarried) women, considered together. Whatever the direct impact of subsidizing fertility among the unmarried, it must be more than offset by the indirect effect of welfare programs reducing the prevalence of intact marriages and thereby reducing fertility. It remains unclear how much of the negative effect of the U.S. welfare benefits on marriage is due to the safety net provided to all poor mothers, regardless of their marital status, and how much is due to married mothers being ineligible for most of these programs. The latter restriction of child allowances to unmarried mothers could be eliminated from the U.S. welfare system, if the support for children in poor families did not discriminate against married women. This change would make the U.S. welfare system more similar to that in Canada and many other industrially advanced nations (Smeeding et al., 1988). Medicaid reimbursements appear to have a less negative effect on marriage in 1990 than in 1980, which could be a reflection of recent welfare and health legislation in the United States that has sought to increase access for the poor to prenatal and child health care, regardless of whether the woman qualified as an unmarried mothers for AFDC benefits (Yelowitz, 1994).

Another change from 1980 to 1990 is for the negative coefficient on women's wages in the fertility equation to have decreased in absolute value while the positive coefficient on men's wages have increased. This pattern could be explained if women's time is becoming a smaller share of private child care costs, and market-provided child care services have correspondingly increased (Schultz, 1981). In other words, as the family searches for lower-cost child care arrangements, the female-time intensity of children is declining. The increasing positive elasticity of fertility with respect to male wages does not signal that the male-time intensity of children is taking up the slack (Schultz, 1981). Whether this increase in child inputs purchased from the market will halt the decline in the prevalence of marriage is a question warranting further study. Some European countries have provided extensive public subsidies for child care as in Sweden, and for preschool programs in France to promote the mother's reentry into the labor force. U.S. welfare reforms initiated in 1996 may also have the side effect of increasing public subsidies for child care for poor mothers, as these women seek employment after exhausting their now rationed welfare benefits. It is time to evaluate how the increased use of market-provided child care is affecting various dimensions of child quality, recognizing that marriage, fertility, and child care arrangements are jointly determined with child quality.

A third change noted from 1980 to 1990 is the increased positive weight of the prospective husband's wage on the likelihood that women will be currently married. Are women assigning a higher value to marrying higher wage-earning men? Are men willing to transfer more of the economic gains they obtain in the labor market to their prospective wives? More structured bargaining models of the family may help us clarify in the future what this increased sensitivity of the marriage market to men's wage opportunities implies for the intrafamily allocation of well-being (McElroy, 1990).

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TABLE A-1

PROBIT COEFFICIENT ESTIMATES FOR THE PROBABILITY OF BEING IN THE WAGE EARNER SAMPLE AND THE LOG WAGE EQUATION FOR BLACK AND WHITE WOMEN AGE 15-65: JOINT MAXIMUM LIKELIHOOD ESTIMATES[a]

Explanatory variables	For a black woman: Prob. of being a wage earner	For a black woman: Log of hourly wage	For a white woman: Prob. of being a wage earner	For a white woman: Log of hourly wage	For the Potential Husband of a black woman: Prob. of being a wage earner	For the Potential Husband of a black woman: Log of hourly wage	For the Potential Husband of a white woman: Prob. of being a wage earner	For the Potential Husband of a white woman: Log of hourly wage
Years of schooling:								
Primary	-.0155 (.86)	-.0199 (1.48)	-.0598 (3.34)	-.0148 (.92)	-.0273 (1.41)	-.0172 (1.24)	.0066 (.36)	.0332 (2.52)
Secondary	.217 (13.3)	.168 (14.0)	.136 (8.03)	.104 (8.26)	.0598 (3.41)	.0974 (6.19)	.0390 (2.27)	.115 (8.49)
Higher	.122 (10.2)	.161 (23.1)	.0529 (5.99)	.131 (27.3)	.0551 (5.02)	.0910 (8.99)	-.0362 (4.52)	.0636 (10.3)
Potential Experience	.0289 (6.34)	.0431 (14.5)	.0056 (1.20)	.0280 (10.4)	.0305 (6.61)	.0320 (8.24)	.0358 (8.57)	.0504 (15.1)
Potential Experience Squared ($\times 10^{-2}$)	-.0737 (7.89)	-.0752 (11.7)	-.0516 (5.54)	-.0485 (8.09)	-.0760 (7.65)	-.0600 (7.20)	-.111 (12.7)	-.101 (14.1)
Hispanic (= 1)	.0481 (.44)	.0679 (.87)	-.0961 (1.62)	-.0278 (.71)	.490 (4.63)	.162 (1.98)	-.112 (.19)	-.101 (2.15)
Woman's Property Income (\$ per year $x10^{-3}$)	.0220 (.51)		-.0194 (2.11)		.0660 (.65)		-.0278 (3.40)	
Woman's Property Income Squared ($\times 10^{-6}$)	.0183 (.00)		.511 (.27)		-37.4 (1.49)		.0238 (1.14)	
Duration of Unemployment Benefits (weeks)	-.0215 (2.91)		-.0162 (2.51)		-.0162 (2.25)		-.0012 (.22)	
AFDC Benefit (\$ per month $\times 10^{-2}$)	-.0806 (5.33)		-.0278 (2.07)		-.0531 (3.61)		-.0511 (4.55)	
Medicaid Reimbursement (\$ per month $x10^{-2}$)	-.0216 (.59)		-.0346 (1.23)		-.0309 (1.03)		-.0526 (2.24)	
Intercept	.385 (1.78)	.745 (7.27)	1.04 (5.22)	1.23 (10.4)	-.100 (.46)	1.17 (8.70)	.178 (.184)	.971 (9.67)
Sigma/Rho	.706 (82.2)	.662 (26.6)	.647 (75.0)	.350 (6.04)	.757 (32.1)	.715 (20.6)	.781 (69.8)	.740 (57.0)
-Log Likelihood	8102.7		9997.6		6043.5		9583.6	
Mean of Dependent Variable (SD)	.675 (.468)	2.01 (.695)	.673 (.469)	2.08 (.688)	.322 (.468)	2.36 (.628)	.531 (.499)	2.56 (.677)
Sample Size	6660	4497	8123	5464	6600	2150	8123	4311

a The absolute value of the asymptotic t ratio is reported in parentheses below each coefficient. The numbers reported in the row labeled "sigma/rho" include first the standard error of the probit equations followed by the correlation of the errors from the probit wage earner and log-linear wage rate equations that are here estimated by joint maximum likelihood methods. Beneath sigma and rho in parentheses is the ratio of this estimate to its standard error.