

## Why Do Married Men Earn More than Unmarried Men?

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Previous research reported that married men, *ceteris paribus*, earn more than unmarried men. A variety of explanations have been suggested for this association. One class of explanation argues that wives, for various reasons, increase their husbands' wages. Another class of explanation maintains that the causality is reversed, that is, high wage men are more likely to get married than low income men. Yet a third possibility is that unobserved characteristics, affecting both wages and marital status, are the reason for the observed cross-sectional association between marital status and wages. This paper presents a longitudinal model suitable for testing these explanations. The use of the model is illustrated by analyzing the wages and marital status of a large sample of men drawn from the Panel Study of Income Dynamics. Significant cross-sectional effects of marital status on wages and vice versa disappeared when the longitudinal model was employed. This suggests that omitted variables affecting both wages and marital status, rather than the former explanations, are responsible for the higher wages of married men. © 1991 Academic Press, Inc.

### THEORETICAL BACKGROUND

That married men earn more than unmarried men is an established finding of many cross-sectional income and wage determination studies. The large positive wage effect of marriage for men persists even in the presence of numerous productivity control variables including hours worked and self-imposed restrictions on where and when to work (Kallechek and Raines, 1976; Hill, 1979; Pfeffer and Ross, 1982). However, as is the case with other characteristics determining wages (e.g., education), a variety of explanations have been suggested for the positive relationship between men's wages and marital status. It is useful to classify these various explanations according to whether they focus on

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wives' effects on men's wages in the labor market or on matching processes in the marriage market.

The more prevalent class of explanation focuses on *wives' effect* on their husbands' wages in the *labor market*. It includes a variety of reasons for employers to grant married men wage premiums. Some maintain that employers respond to an actual productivity increase caused by the existence of a wife. Depending upon the writer, wives are said to improve the household's decision making process, to motivate men to put more effort into their jobs, to provide emotional support and advice on job-related matters, and to perform tasks directly related to their husbands' jobs (Benham, 1974; Kanter, 1977; Pfeffer and Ross, 1982).<sup>1</sup> Others draw on theories of statistical discrimination and market signaling (Spence, 1974), suggesting that employers *perceive* married men to be more stable, responsible, and hence more productive than similar unmarried men (Block and Kuskin, 1978). Yet others have raised the possibility that employers, conforming to dominant values and norms, discriminate against unmarried men (Talbert and Bose, 1977) or perceive them to be in lesser financial need than their married counterparts (Pfeffer and Ross, 1982; Hill, 1979).

Notwithstanding the many important differences among these explanations, they all share the belief that the existence of a wife increases a man's wage, whatever the reason. In other words, all the explanations discussed thus far predict an increase in a man's wage rate following his marriage, and a decrease following divorce, separation, or death of a wife. We will henceforth refer to this prediction as the "wives' effect" hypothesis.

This last prediction (decrease in wages following divorce or death of a wife) is particularly true for the explanations focusing on wives' effect on men's actual productivity (e.g., Kanter, 1977; Benham, 1974). According to these explanations, when there is no longer a wife enhancing her husband's productivity, the man's on-the-job performance and hence wages will decline. The other variants of the "wives' effect" hypothesis are less clear on that matter. The need explanation (Pfeffer and Ross, 1982) has no clear prediction on the effect of divorce on men's wages, for such men are often perceived as being in great financial needs. Likewise, the signaling hypothesis would not expect divorce to depress a man's wage rate if the employer has already gathered enough information on the man's performance. In such case, the employer will no longer be

<sup>1</sup> Human capital theory, too, believes that marriage increases men's productivity. Rarely, however, have human capital students elaborated on the mechanisms by which wives increase their husbands' productivity. In most studies (e.g., Kalechek and Raines, 1976; Oaxaca, 1973; Polachek, 1975) marital status was included as a component of human capital in wage models for men, with no rationale being furnished for its inclusion.

forced to rely on marital status as a performance signal for this particular worker. This being the case, the "wives' effect" hypothesis as formulated above is more appropriate for testing theories stressing the impact of wives on men's actual on-the-job performance than on theories focusing on employers perceptions of workers needs or performance.

The second type of explanation for the positive association found in cross-sectional data between men's wages and being married focuses on matching processes in marriage markets, thus reversing the causal order. As noted by Gwartney and Stroup (1973, p. 585) "Males who remain single might do so because most females accurately perceive that they will be unlikely to attain economic success." In other words, high income men may be more attractive in marriage markets and therefore more likely to get married than other men. It is thus possible that men's wages affect their propensity to get married and divorce rather than vice versa. We henceforth refer to this approach as the "marriage market" hypothesis.

To be sure, the two hypotheses (and the various explanations) discussed above are not mutually exclusive. Thus, it is possible that the observed relation between men's wages and their marital status found in numerous cross-sectional studies is in part due to a change in men's wages following marriage, and in part due to the higher propensity of high wage men to get married (cf. Becker, 1981, p. 67). Previous empirical research, however, focused solely on the "wives' effect" hypothesis, and did not test the reverse "marriage market" hypothesis. Moreover, it relied on cross-sectional models for testing this hypothesis (Gwartney and Stroup, 1973; Oaxaca, 1973; Rosensweig and Morgan, 1976; Hill, 1979; Pfeffer and Ross, 1982).<sup>2</sup> Such models, as we will demonstrate below, cannot overcome the problem of omitted variables (see Pindyck and Rubinfeld, 1981, pp. 128-130 for a discussion of the problem of omitted variables), and are therefore inappropriate for testing these hypotheses.

Our purpose in this paper is to develop a longitudinal model suitable for testing both the "wives' effect" and the "marriage market" hypotheses. The next section will discuss the problem of omitted variables in cross-sectional models aimed at analyzing the relationship between marital status and wages, and will present an appropriate longitudinal model for evaluating the empirical status of the above hypotheses. Next, we will illustrate the advantages of the proposed model using data drawn from the Panel Study of Income Dynamics. Finally, we discuss these results and their possible implications for the various theoretical explanations.

<sup>2</sup> Bartlett and Callahan's longitudinal study (1984) is an exception discussed in footnote 6.

## THE MODEL

The conventional cross-sectional model for estimating the "wives' effect" hypothesis is

$$\ln W_t = X_t b_t + a_t M_t + e_t, \quad (1)$$

where the subscript  $t$  denotes time,  $W$  represents wage,  $X$  is a vector of wage covariates,  $b$  denotes a vector of their coefficients,  $M$  is marital status,  $a$  is its coefficient, and  $e$  is the normally distributed, independent, error term.

In order to measure the effect of marital status on wages at a second point in time ( $t + 1$ ), we use the exact same model as Eq. (1):

$$\ln W_{t+1} = X_{t+1} b_{t+1} + a_{t+1} M_{t+1} + e_{t+1}. \quad (2)$$

These cross-sectional models are unable to deal with the problem of omitted variables. It is therefore possible that unobserved characteristics affect *both* men's propensity to be married *and* to earn high wages. Conformity to social expectations is an example of a possible omitted variable. It is reasonable to argue that men who score low on this unobserved variable may be less concerned about satisfying social expectations by remaining single and would also be more likely to defy the norm of striving to get ahead socially and economically. To the extent that such characteristics are omitted from the analyses, the error terms  $e_t$  and  $e_{t+1}$  are no longer random. They include the effects of the unobserved variables on wages, and are correlated with the marital statuses  $M_t$  and  $M_{t+1}$ , thus causing the coefficient of marital status to "capture" the effects of the omitted variables on wages, and leading to an incorrect estimate of marital status on men's earnings.

Now consider a longitudinal model where the dependent variable is a *wage change* from time  $t$  to time  $t + 1$ . Subtracting Eq. (1) from Eq. (2) (first-differencing) yields the desired model

$$\begin{aligned} \ln W_{t+1} - \ln W_t = & (X_{t+1} - X_t) b_{t+1} + X_t (b_{t+1} - b_t) + a_{t+1} (M_{t+1} - M_t) \\ & + (a_{t+1} - a_t) M_t + (e_{t+1} - e_t). \end{aligned} \quad (3)$$

The coefficient of interest is the one for marital status change from time  $t$  to time  $t + 1$  ( $a_{t+1}$ ). If there is a marital status effect on wages (i.e., a wage increase following marriage, and a decrease following divorce, separation, or death of wife), this coefficient should be positive when associated with a status change from unmarried to married, and negative following a change from a married to an unmarried status. Close

examination of the wage change model indicates not only that the two effects are opposite in sign but are necessarily also equal in magnitude.<sup>3</sup>

The new error term ( $e_{t+1} - e_t$ ) is random and overcomes the problem of time invariant omitted variables.<sup>4</sup> If unobserved characteristics which are correlated both with marital status and with wages are included in  $e_t$ , they must also be part of  $e_{t+1}$ , as the cross-sectional models [Eqs. (1) and (2)] refer to the *same* men at two points in time. By subtracting  $e_t$  from  $e_{t+1}$  we eliminate these unobserved characteristics of those changing their marital status from the wage change model.<sup>5</sup> Thus, a possible problem of omitted variables in the cross-sectional model is controlled for in this longitudinal model by examining the two groups experiencing marital status changes.<sup>6</sup>

Turning now to the "marriage market" hypothesis, the cross-sectional model for estimating earning effect on marital status is

$$M_t = X_t c_t + d_t \ln W_t + u_t, \quad (4)$$

where the subscript  $t$  denotes time,  $M$  is a categorical variable indicating whether a person is married or not,  $X$  is a vector of marital status correlates,  $c$  denotes a vector of their coefficients,  $W$  is wage,  $d$  is its coefficient, and  $u$  is the error term. If we refer to the same model at a second point ( $t + 1$ ) in time we obtain:

$$M_{t+1} = X_{t+1} c_{t+1} + d_{t+1} \ln W_{t+1} + u_{t+1}. \quad (5)$$

Here again, these cross-sectional models cannot deal with the problem

<sup>3</sup> This constraint is part of the marital status change expression  $a_{t+1}(M_{t+1} - M_t)$ , because the coefficient  $a_{t+1}$  is the *same* when changing marital status from married to unmarried and vice versa. The only difference between the two is the sign of this coefficient. For another example of this constraint see Mellow (1981) who estimated a similar longitudinal model for testing unionization effect on wages.

<sup>4</sup> It does not overcome the problem of omitted variables which are related to changes in marital status and whose values change over time. Such omitted variables, however, are unlikely in our case.

<sup>5</sup> It is thus apparent that the potential problem of omitted variables *cannot* be overcome using the alternative longitudinal model prevalent in the literature (Hannan, 1979):

$$\ln W_{t+1} = \ln W_t + X_t b_t + a_t M_t + e_{t+1}. \quad (3a)$$

<sup>6</sup> Bartlett and Callahan (1984) used a longitudinal model for estimating the effect of changes in marital status on wages, but their model suffers of several problems. First, their results fail to satisfy the constraint discussed in footnote 3. This may be due, at least in part, to the fact that they include in their wage change model men who have been married continuously from the first to the second points in time. Furthermore, their sample (NLS of white older men) is not representative of the marriage market. Finally, they measured a wage change between 1976 to 1977, and marital status change between 1966 to 1977. Although the time lag for wives' effect on their husbands' wages (if any) is unknown, it is difficult to believe that a wage change in a certain year could be the result of a change in marital status occurring as many as 10 years previously.

of omitted variables. A longitudinal model of marital status change is necessary for overcoming this problem:

$$M_{t+1} - M_t = (X_{t+1} - X_t)c_{t+1} + X_t(c_{t+1} - c_t) + d_{t+1}(\ln W_{t+1} - \ln W_t) + (d_{t+1} - d_t)\ln W_t + (u_{t+1} - u_t). \quad (6)$$

If both coefficients—that of marital status change in Eq. (3) ( $a_{t+1}$ ) and that of wage change in Eq. (6) ( $d_{t+1}$ )—are significant, it implies that both the “wives effect” and “marriage market” hypotheses must not be rejected.<sup>7</sup>

### EMPIRICAL ILLUSTRATION

In order to illustrate our argument, we use data drawn from the Panel Study of Income Dynamics (PSID). The PSID has followed the economic fortunes of some 5000 families since 1968, each year interviewing one member of the original sample families. As individuals leave their original families (e.g., divorced persons), the PSID follows them in their new households. These particular longitudinal features of the PSID enable us to select a subset of males suitable for testing the models presented above. The specific subset to be used in the analyses below includes all males aged between 24 and 64 years in 1983 for whom marital status and labor market information is available for the years 1979–1982.<sup>8</sup> There are 3215 such males.<sup>9</sup>

Table 1 describes and presents the means and standard deviations of the variables used in the analyses. The variables of interest are those indicating changes. The difference between wages earned in 1982 and 1979 is measured by CHGWAGE, and the difference between wages earned in 1981 and 1979 is measured by CHGWAGE1. Changes in marital status between 1979 and 1982 are indicated by CHGTOMAR (an un-

<sup>7</sup> As is the case with Eq. (3a), it is apparent that the following alternative longitudinal model,

$$M_{t+1} = M_t + X_t c_t + d_t \ln W_t + u_{t+1}, \quad (6a)$$

does not overcome the possible problem of omitted variables.

<sup>8</sup> Each year the PSID collects labor force information, including components of income for the preceding year. Thus, 1983 interviews include wage and labor force information for 1982.

<sup>9</sup> Since complete labor market and marital status information is available for heads of household only, our sample is practically composed of all male heads of household, 24–64 years of age in 1983 who have been heads of household for at least 4 years. To be sure, men who are heads of household are more likely than other men to be *both* married *and* of high wages. Therefore, using such a sample in cross-sectional analysis would result in sample selection bias [see Berk (1983) for a discussion of this topic]. But, the longitudinal model to be used overcomes this problem in the same manner by which it avoids the problem of omitted variables. However, the results can be generalized only to males 24 to 64 years of age.

TABLE 1  
Variables Used in the Analyses: Definitions, Means, and Standard Deviations  
(in Parentheses)

Variable	Definition	Mean (S.D.)
LNHW82	ln (hourly wage in 1982)	1.802 (1.504)
LNHW81	ln (hourly wage in 1981)	1.854 (1.264)
LNHW79	ln (hourly wage in 1979)	1.768 (1.038)
CHGWAGE	LNHW82-LNHW79	.033 (1.264)
CHGWAGE1	LNHW81-LNHW79	.090 (1.022)
MAR82	1 if married in 1982, 0 otherwise	.840 (.367)
MAR79	1 if married in 1979, 0 otherwise	.831 (.374)
CHGTOMAR	1 if not married in 1979 and married in 1982, 0 if not married both years	.230 (.421)
CHGTOSNG	1 if married in 1979 and not married in 1982, 0 if married both years	.036 (.187)
WHITE	1 if white, 0 otherwise	.711 (.453)
MGR	1 if manager or professional, 0 otherwise	.280 (.449)
ED	Years of schooling	12.173 (2.961)
EXP79	Years of labor market experience	10.527 (10.811)
AGE	Years of age	39.952 (11.303)

married man in 1979 who was married by 1982) and CHGTOSNG (a married man in 1979 who was unmarried by 1982). The remaining independent variables—race, occupation, years of schooling, and age or experience<sup>10</sup>—reflect 1979 information only. Such being the case, we are forced to assume that there were no changes in these variables during the time interval investigated.<sup>11</sup>

<sup>10</sup> Because of the high correlation between age and experience, we never use both variables in the same equation. Since labor force experience is more relevant in the labor market than age, we include the former in wage equations (7)–(9); and since the opposite is true for the marriage market, we use age rather than experience in marital equations (10)–(13).

<sup>11</sup> Unfortunately, the PSID does not update education, occupation, and experience in-

TABLE 2  
Means and Standard Deviations (in Parentheses) for the Married and Unmarried Subsamples by Years

Variable	1979		1982	
	Married	Unmarried	Married	Unmarried
LNHW82	1.843 (1.535)	1.579 (1.320)	1.855 (1.525)	1.500 (1.354)
LNHW79	1.819 (1.023)	1.499 (1.077)	1.819 (1.025)	1.478 (1.066)
WHITE	.739 (.439)	.565 (.496)	.743 (.437)	.538 (.499)
MGR	.287 (.453)	.238 (.426)	.291 (.454)	.218 (.413)
ED	12.152 (2.956)	12.229 (2.993)	12.167 (2.952)	12.159 (3.018)
EXP79	11.425 (10.799)	6.026 (9.735)	11.223 (10.786)	6.801 (10.207)
Sample size	2698	547	2726	519
	Continuously married <sup>a</sup>	Continuously unmarried <sup>a</sup>	CHGTOMAR <sup>a</sup>	CHGTOSNG <sup>a</sup>
Mean	.031	.061	.142	-.146
SD	1.303	1.030	1.000	1.309
n	2600	421	126	98

<sup>a</sup> The means and standard deviations of CHGWAGE, by marital status.



TABLE 3  
Pearson Correlations among the Variables Used in the Analyses

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)	(13)
LNHW79	1.0												
LNHW81	.62	1.0											
LNHW82	.56	.62	1.0										
CHGWAGE	-.16	.23	.73	1.0									
CHGWAGE1	-.25	.60	.20	.44	1.0								
CHGTOMAR	-.11	.11	.11	.03	.03	1.0							
CHGTOSNG	-.03	-.03	-.04	-.03	-.01	—	1.0						
MAR79	.12	.10	.07	-.02	.00	—	—	1.0					
MAR82	.12	.11	.09	.00	.01	1.0	-1.0	.75	1.0				
WHITE	.21	.19	.18	.04	.02	.12	-.08	.14	.17	1.0			
MGR	.26	.26	.25	.08	.06	.00	-.07	.04	.06	.27	1.0		
ED	.31	.31	.31	.11	.07	.05	.00	-.01	.00	.31	.50	1.0	
EXP79	-.01	-.10	-.12	-.13	-.11	-.13	-.08	.19	.15	.06	-.03	-.26	1.0
AGE	-.04	-.12	-.14	-.13	-.11	-.13	-.07	.19	.14	.06	-.01	-.25	-.97

TABLE 4  
Wage Regressions (*t* Values Are in Parentheses)

Variable	LNHW79 (1)	LNHW82 (2)	CHCWAGE (3)	CHGWAGE (4)
WHITE	.202* (4.96)	.232* (3.89)	-.007 (-.11)	.245** (2.41)
MGR	.270* (6.07)	.388* (5.99)	.091 (1.39)	.252** (2.11)
ED	.081* (11.34)	.106* (10.20)	.030* (2.85)	-.002 (-.10)
EXP79	.003 (1.74)	-.011* (-4.54)	-.015* (-6.16)	-.012* (-2.57)
MAR79	.261* (5.55)			
MAR82		.317* (4.56)		
CHGTOMAR				-.086 (-1.32)
CHGTOSNG			-.01 (-.62)	
CONSTANT	.316 (3.39)	.089 (.65)	-.152 (-1.09)	.341 (.98)
<i>R</i> <sup>2</sup>	.130	.121	.028	.040
<i>N</i>	3215	3215	2673 <sup>a</sup>	542 <sup>b</sup>

<sup>a</sup> Sample includes only those who remained married and those who changed their status from married to single.

<sup>b</sup> Sample includes only those who remained single and those who changed their status from single to married.

\*  $p < .01$ .

\*\*  $p < .05$ .

Table 2 presents descriptive statistics for the married and unmarried men, and Table 3 details the zero order correlations among the variables used in the analyses. The only variable for which both groups have the same mean is education. On all other variables, married men score, on the average, higher than unmarried men. All these differences are significantly different from zero ( $p < .001$ ). This pattern holds true for both 1979 and 1982. Thus, married men have similar education as their unmarried counterparts, but the former are more likely to be whites, profes-

formation every year. Rather, data values for these variables were carried forward from 1976 (or, in cases of men who became heads after 1976, from the year they became heads of household (Survey Research Center, 1985). However, our interest in this paper is limited to marital status and wage variables only. Thus, education and experience are included as controls for human capital, occupation (measured by a dummy variable for the top two occupational categories: managers and administrators and professional and technical workers) as a crude proxy for job characteristics, and race for its known large effect on wages.

sionals or managers, with more labor force experience, and of higher wages than the latter.

We now proceed to estimate seven equations, as follows: First, we use OLS regressions for two cross-sectional wage functions, one for 1979, and one for 1982 [corresponding to Eqs. (1) and (2)].

$$\text{LNHW79} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{EXP79}, \text{MAR79}) \quad (7)$$

$$\text{LNHW82} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{EXP79}, \text{MAR82}). \quad (8)$$

The first two columns of Table 4 present the results for these cross-sectional regressions. The coefficients of interest are those of MAR79 and MAR82. Since they are positive and statistically significant we conclude that in the population represented by the PSID (as in most other studies) married men indeed earn more than unmarried men when all other measured variables are held constant. Next, we estimate the longitudinal wage equation [corresponding to Eq. (3)] where the dependent variable is wage change from 1979 to 1982:

$$\text{CHGWAGE} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{EXP79}, \text{MAR79}, \quad (9) \\ \text{CHGTOMAR}, \text{CHGTOSNG}).$$

This equation is estimated separately for those who changed their status from married to single, and for those who changed their status from single to married.<sup>12</sup> Here we are interested primarily in the coefficients of CHGTOSNG in Eq. (3) and CHGTOMAR in Eq. (4).<sup>13</sup> As shown above, if there is a marital status effect on wages (the "wives' effect" hypothesis), these two coefficients should be statistically significant, similar in magnitude, and opposite in sign (CHGTOMAR should be positive). The results of this change model (columns 3 and 4 of Table 4) reveal that the coefficient of CHGTOSNG is much smaller than that of CHGTOMAR, and both carry the same sign. A change from married in 1979 to unmarried in 1982 is associated with a 1% reduction in wages; the opposite change is associated with a reduction of 8.6%. Notwith-

<sup>12</sup> We decided to estimate two separate equations because the correlation between CHGTOMAR and CHGTOSNG is  $-.96$ . Estimating the model including both variables in one equation would produce extremely unstable coefficients.

<sup>13</sup> Due to possible simultaneous effect of changes in marital status on changes in wages and vice versa, predicted rather than observed values of CHGTOMAR and CHGTOSNG are used in Eqs. (3) and (4). These values were derived using an instrument which was found to be correlated with CHGTOMAR ( $r = .53$ ) and with CHGTOSNG ( $r = -.21$ ), and uncorrelated with CHGWAGE ( $r = .00$ ). This instrument is 1983 wife's annual hours of housework, and it was identified empirically. While admittedly problematic, empirical identification of instruments is still more useful for our purpose than no identification at all. The other variables used to identify CHGTOMAR and CHGTOSNG were WHITE, MGR, ED, LNHW79, and AGE.

standing this difference both coefficients of marital status change (from unmarried to married and from married to unmarried) are not significantly different from zero, and therefore do not violate the model's constraint. Thus, although there was a significant wife effect in the cross-sectional models, it disappeared when the effects of omitted variables on marital status change were controlled by means of a longitudinal analysis. This finding is inconsistent with the "wives' effect" hypothesis common to the variety of explanations suggesting that men's wages are affected by changes in their marital status.<sup>14</sup>

The "marriage market" hypothesis is tested using logit analyses. Again, we begin with two cross-sectional marital status equations, one for 1979, and one for 1982 [corresponding to Eqs. (4) and (5), respectively]:

$$\text{MAR79} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{AGE}, \text{LNHW79}). \quad (10)$$

$$\text{MAR82} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{AGE}, \text{LNHW82}). \quad (11)$$

If there is a wage effect on marital status, we should observe positive coefficients for LNHW79 and LNHW82. Table 5 presents the logistic regressions for these models. The cross-sectional analyses yield positive and statistically significant wage effects on marital status both in 1979 and in 1982 (columns 1 and 2). Next, we estimate the longitudinal marital status change model [corresponding to Eq. (6)] from 1979 to 1981. This is done by estimating two equations: one for status change from married to unmarried; the other for the opposite change:<sup>15</sup>

$$\text{CHGTOSNG} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{AGE}, \text{LNHW79}, \text{CHGWAGE1}). \quad (12)$$

$$\text{CHGTOMAR} = f(\text{WHITE}, \text{MGR}, \text{ED}, \text{AGE}, \text{LNHW79}, \text{CHGWAGE1}). \quad (13)$$

<sup>14</sup> Using the same data we also tested the alternative model below (9a) based on Eq. (3a). The results of this model (9a) regarding the effects of marital status on wages were appreciably the same to those obtained using the CHGWAGE Eq. (9); namely, no significant effect of marital status on wages was detected (*t* values in parentheses,  $R^2 = .343$ ):

$$\text{LNHW82} = -.104 + .734\text{LNHW79}(32.99) + .093\text{WHITE}(1.81) + .195\text{MGR}(3.46) + .046\text{ED}(5.02) - .014\text{EXP}(-6.35) + .085\text{MAR79}(1.43). \quad (9a)$$

<sup>15</sup> Using a multinomial model rather than two separate logit equations is inappropriate here because each person included in the sample can take one out of *two* values only, and *not* one out of four. The four possible categories of the dependent variable are married in both 1979 and 1982 (married-married); single in both 1979 and 1982 (single-single); married in 1979 and single in 1982 (married-single); and single in 1979 and married in 1982 (single-married). It is obvious that a man's marital status in 1979 determines two categories only into which he can be classified (e.g., a man who was married in 1979 can be in either the married-married or married-single categories, but not in the single-single or single-married categories). Since subjects cannot assume all four possible categories, the analyses should be confined to those two possibilities which they can have. Such analyses require two separate equations.

TABLE 5  
Logit Analyses (*t* Values Are in Parentheses)

Variable	MAR79 (1)	MAR82 (2)	CHGTOSNG (3)	CHGTOMAR (4)
WHITE	.346* (6.45)	.426* (7.90)	-.355* (-3.21)	.307** (2.55)
MGR	.005 (.07)	.090 (1.31)	-.509* (-3.01)	-.119 (-.83)
ED	-.018 (-1.57)	-.025** (-2.23)	.047 (1.91)	-.031 (-1.26)
AGE	.024* (9.32)	.018* (7.15)	-.016* (-2.86)	-.019* (-3.17)
LNHW79	.130* (5.58)	—	-.056 (-1.18)	.191** (2.56)
LNHW82	—	.073* (4.86)	—	—
CHGWAGE1			-.048 (-1.09)	-.084 (-1.27)
Constant	4.646	5.015	3.841	4.950
<i>N</i>	3232	3232	2687 <sup>a</sup>	545 <sup>b</sup>

<sup>a</sup> Sample includes only those who remained married and those who changed their status from married to single.

<sup>b</sup> Sample includes only those who remained single and those who changed their status from single to married.

\*  $p < .01$ .

\*\*  $p < .05$ .

The focal coefficient here is that of CHGWAGE1.<sup>16</sup> If wages affect marital status, then we expect this coefficient to be negative in Eq. (12), and positive in Eq. (13). However, the results obtained suggest otherwise: the wage change between 1979 and 1981 did not bring about a change in marital status from married in 1979 to unmarried in 1982 (column 3, Table 5) nor a change from unmarried in 1979 to married during the same time period (column 4, Table 5). In short, the longitudinal regressions lend no support to the “marriage market” hypothesis.<sup>17</sup>

<sup>16</sup> Recall that CHGWAGE1 is the wage change between 1979 and 1981, rather than 1982. We use it here to allow a time lag between a change in wages and a change in marital status.

<sup>17</sup> As with the wage equation, we also tested the alternative model (13a) below, based on model (6a). Logistic analysis of this model, however, yielded different results from those presented in Table 5. Here we observe a significant effect of wage in 1979 on marital status in 1982, even when marital status in 1979 is held constant (*t* values in parentheses):

$$\text{MAR82} = 4.46 + .361\text{WHITE}(4.39) + .170\text{MGR}(1.70) - .034\text{ED}(-2.15) \\ - .001\text{AGE}(-.19) + 2.21\text{MAR79}(29.26) + .069\text{LNHW79}(2.150). \quad (13a)$$

## DISCUSSION AND CONCLUSIONS

The results presented in Tables 4 and 5 indicate that the observed cross-sectional association between wages and marital status found in the data is due neither to an increase in men's wages following marriage nor to a higher propensity to marriage among economically successful men. We base this conclusion on the fact that cross section effects of marital status on wages and vice versa vanish when a longitudinal analysis is used.

In a recent study, Nakosteen and Zimmer (1987) failed to detect an effect for marital status on wages estimating a cross-sectional simultaneous equations model using a sample of 567 men 18–24 years old. At first stage, they estimated a marital status equation, with parents' education, number of siblings, presence of older siblings, religion, and urban upbringing serving as instrumental variables. A significant effect for marital status that was found in a conventional wage model disappeared when the two-stage model was estimated.

The present study offers a different and more comprehensive examination of the problem. First, it utilizes a longitudinal rather than cross-sectional model. Doing this it avoids the nearly impossible task of theoretically identifying instrumental variables for a marital status equation. Indeed, the family background variables used by Nakosteen and Zimmer (1987) to identify married status (e.g., religion, number of siblings, urban upbringing) have been found by numerous studies to affect income, (Featherman and Hauser, 1978; Jencks *et al.*, 1979; Cohen, 1986). Second, our study does not restrict itself to processes occurring in the labor market only. It pays equal attention to the marriage market; the determination of marital status is studied, and does not serve merely as a first stage estimate for a wage equation. Finally, the sample used here is larger, and more importantly is not restricted to young men. This enabled us to examine young as well as older males and to study the effects and determination not only of marital status change from unmarried to married but also from married to unmarried.

Should future research, using different and perhaps better data, obtain results similar to ours, we would have more faith in the omitted variables explanation. Better yet, of course, is to identify these characteristics. We have speculated above that one such unobserved characteristic may be conformity to social expectations. Another possibility is that good looking men are more attractive in both marriage and labor markets. There is evidence that physical features of men affect their success in some bureaucracies (Mazur, Mazur, and Keating, 1984), and it is reasonable to expect similar features (e.g., height) to affect their propensity to marry.

It is also possible that these omitted variables are social structural

rather than individual. Consider, for example, men's positions within social networks. Both wives and high-paying jobs are known to be found via social networks. Hence, men who are well placed within the structure of social relationships are more likely to be both married and of higher wages than other men. To be sure, these are not the only possible unobserved characteristics affecting both wages and marital status. Future research, failing to detect wives or marriage market effects in longitudinal models, should try to identify such characteristics.

A related issue to be explored by future research is the cross-sectional results revealing higher wages among unmarried women than among married women. To be sure, the reasons for these cross-sectional findings may be due to labor market processes, and/or marriage market processes and/or omitted variables affecting both women's propensity to be unmarried and to command high wages. That our conclusion regarding men focuses on omitted variables as the key explanation does not necessarily imply that this is the explanation for the wage premium enjoyed by unmarried women compared to married women. It is possible that among women omitted variables affecting wages and the propensity to be married are only secondary in importance to effects of marital status on wages and/or effects of wages on marital status. It is still an open empirical question.

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