Does Marriage Really Make Men More Productive?

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ABSTRACT

This paper presents new descriptive evidence regarding marital pay premiums earned by white males. Longitudinal data indicate that wages rise after marriage, and that cross-sectional marriage premiums appear to result from a steepening of the earnings profile. Data from a company personnel file that includes information on job grades and supervisor performance ratings reveal large marital status pay differences within a narrow range of occupations (managers and professionals) and environments (a single firm). Married workers tend to be located in higher paying job grades; there are very small pay differentials within grades. Married men receive higher performance ratings than single men; as a result, they are much more likely to be promoted. Controlling for rated performance, however, eliminates the promotion differential.

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I. Introduction

Labor economists have long noted that married men earn substantially more per hour worked than men who are not currently married. These cross-sectional wage differentials persist when controls are introduced for education, race, region, age, or work experience, and even occupation and industry. Typically, differentials are in the 10–40 percent range—roughly as large as race, firm-size, and union wage differentials, as well as differentials across industries, all of which have been extensively studied. The marriage premium is of particular interest for estimating gender-based discrimination in labor markets, as the male marital pay premium accounts for about one-third of estimated gender-based wage discrimination in the United States (e.g., Neumark 1988).[1] More generally, efforts to explain the marriage wage premium can contribute to achieving a better understanding of the determination of individual wages. Marital status differentials in labor market outcomes may also be of increasing interest in light of trends toward delaying (and perhaps foregoing) entry into first marriage, increased divorce rates, and sharply higher labor force participation rates among married women (especially those with young children). If marital status pay differentials reflect productivity differences, then changes in the marital status composition of the labor force potentially can affect the productivity of the labor force.

While there is widespread agreement that cross-sectional marriage differentials for men are sizable, there is much less agreement about their source. In fact, in a recent review of gender wage differentials, Goldin (1990, p. 102) concluded that ". . . the role of marriage in enhancing the earnings of male workers is still only dimly understood." One major hypothesis is that earnings differentials result from productivity differentials: that is, marriage per se makes workers more productive (Becker 1981, 1985; Kenny 1983; Greenhalgh 1980). Another hypothesis attributes these differentials to employer favoritism (Hill 1979; Bartlett and Callahan 1984), and a third to selection into marriage on the basis of wages or personal characteristics that are valued in labor markets (Becker 1981; Nakosteen and Zimmer 1987; Keeley 1977).[2]

1. The standard technique for estimating wage discrimination (Oaxaca 1973, Blinder 1973) decomposes a wage differential between two groups into a part due to differences in average characteristics, and a part due to different coefficients in separate wage regressions. The part resulting from differences in coefficients is commonly interpreted as discriminatory. Marital status coefficients for men are large and positive, while for those for women are typically close to zero.

2. Our review of published research in this area indicates that these categories, broadly defined, cover the majority of hypotheses that have been advanced to date. Additional

This paper presents new descriptive evidence that permits appraisal and refinement of hypotheses regarding the source of marriage pay premiums for white males.[3] The evidence is drawn from two sources: the National Longitudinal Survey of Young Men (NLSYM), and the company personnel file of a single large U.S. manufacturing firm. The longitudinal data are used to examine how wages of individuals change as they change marital statuses, and to examine how wages change for individuals with each additional year spent in a particular marital state. The company personnel data allow us to examine marital status differences within a fairly homogeneous set of occupations (managers and professionals) and environments (i.e., within a single firm), thereby controlling implicitly for important characteristics of workers and jobs that can potentially vary across persons of different marital statuses. These data also include information unavailable from more standard labor market data sources, such as an employee's position (or job grade) in the company, and the performance rating he is given by his supervisor.

In the following section we review the existing empirical literature on male marriage pay differentials. Section III describes the NLS data and Section IV describes the estimation procedures and results. Section V describes the company personnel data and presents empirical findings. A discussion and concluding remarks follow in Section VI.

II. Empirical Studies of the Marriage Wage Differential

That married men earn more than otherwise comparable single men is among the most robust findings from human capital wage equations. A comprehensive review of the multitude of studies that have estimated cross-sectional wage equations with marital status controls would, therefore, be a monumental task, and is beyond the scope of this paper.[4] To cite only the most recent example, Schoeni (1990) documents that male marital pay differentials are large and statistically significant in each of the twelve industrialized countries that he studies. This section presents a brief and selective review of cross-sectional studies of the

hypotheses exist, however, that do not fall into one or more of these three categories. Reed and Harford (1988) have recently suggested that the marriage premium may be a "compensating differential" needed to bribe married men to work under adverse conditions.

3. Korenman and Neumark (1990, 1988) explore the relationships between marriage and pay for women and minority men, respectively.

4. For partial summaries see Nakosteen and Zimmer (1987), Kenny (1983), and Bartlett and Callahan (1984).

marriage wage premium, and a more detailed review of research that explicitly addresses hypotheses regarding the sources of these premiums.

A. *First-Generation Studies (Cross-Sectional Wage Studies)*

Hill (1979), using data from the 1976 Panel Study of Income Dynamics, finds that married men have higher wages than widowed, divorced, or separated men who, in turn, have higher mean wages than never-married men. Her regression analyses include very detailed controls for human capital, work history, health status, occupation, industry, and number of children. In (log) hourly wage regressions for white men, the estimated coefficient of the dummy variable for currently married men is 0.29 with a standard error of 0.04, while it is 0.27 with a standard error of 0.06 for those who are widowed, divorced, or separated.

Greenhalgh (1980) analyzes cross-sections of British workers for the years 1971 and 1975 and finds similar (though somewhat smaller) marital premiums. She attributes the "unexplained" portion of the differentials between married and single men to family role specialization. However, the unexplained portion could also result from other causes such as selection into marriage or employer favoritism.

Bartlett and Callahan (1984) study a sample of men aged 55–64 in 1977, drawn from the National Longitudinal Survey of Older Men, and find that married men earn 20–32 percent more than otherwise comparable unmarried men. Continuously married men, and men who married between 1966 and 1977, had faster wage growth between 1976 and 1977 than other men in the sample. A limitation of this sample of older men is that the effect of first marriages on earnings cannot be estimated; only four men in the sample changed from single to married status in the period covered.

Using wage surveys of three English establishments, Siebert and Sloane (1981) find unadjusted wage differentials favoring married men that range from 15 to 50 percent. Wage equations estimated separately for each of the three establishments (including controls for experience, company tenure, and education) yield differentials of 11, 12, and 27 percent, with corresponding t-statistics of 1.3, 2.0, and 3.9.

B. *Second-Generation Studies*

Two studies use techniques that go beyond the (essentially) cross-sectional analyses reviewed above, and therefore warrant a more detailed review.

Kenny (1983) uses retrospective data for men aged 30–40 in 1969 to compare mean monthly growth rates of earnings in married as compared to not-married months. He regresses the difference between the rate of growth of a man's wages in his married and not-married months, on age growth, and on the rate of growth of the general level of manufacturing wages, in the same months. The positive (and significant) *intercepts* are interpreted as indicating that wage growth is higher in the married months as compared to the not-married months. Adding controls for race, education, age at marriage, verbal ability, and wife's education makes the intercept negative and statistically insignificant. Kenny concludes that higher wage growth rates in the married months indicate that "a very large fraction of the wage differential between married males and never-married males appears to be attributable to the additional investment in human capital that occurs during marriage" (p. 229). Although faster wage growth is consistent with the hypothesis that greater investment in human capital takes place after marriage, Kenny presents no direct evidence on human capital investment. For example, Kenny does not estimate a marriage effect *net* of measured human capital (e.g., experience and tenure), so it is impossible to determine how much of the marital pay differential is attributable to greater accumulation of human capital. Finally, his findings are also consistent with relatively higher "returns" to—rather than greater investment in—human capital in married months that could result from factors such as favoritism, or from changes associated with marriage in tastes, prices, or effort levels.

Nakosteen and Zimmer (1987) correctly point out that "A possibility exists that marital status is determined stochastically by a process whose random unobservable component is correlated with unobservables in the wage/earnings function. In such a case conventional least squares estimates of the wage function, in particular the marital status coefficient and its standard error, are biased and inconsistent" (p. 249). To account for this problem, Nakosteen and Zimmer specify and estimate two-stage models in which endogeneity between marital status and earnings is allowed. The model that receives the most attention in the paper consists of an annual earnings equation for both married and unmarried men, containing a dummy variable for marital status, and a reduced form probit equation for marital status.[5]

5. This is the "treatment effects" model of Barnow, Cain, and Goldberger (1980). The procedure used to obtain consistent estimates of the wage equation parameters is (i) to estimate the probit, (ii) to construct from this probit a "marriage selectivity" variable related to the probability of being married, and (iii) to include this variable on the right hand side of the wage equation, and to estimate this equation by least squares (correcting the standard errors of the coefficient estimates).

Turning to their results, OLS estimation of the (log) annual earnings equation yields a coefficient on the marital status dummy variable of 0.370, with a standard error of 0.186. In the estimates that allow for endogeneity of the form described above, the marital status coefficient rises slightly to 0.410, while its standard error balloons to 0.695. Furthermore, the estimated coefficient of the selectivity variable is insignificantly different from zero, negative, and imprecisely estimated. (The point estimate is -0.190, with a standard error of 4.75.) The authors conclude, despite the large effect of marital status on earnings implied by their point estimates, that "the significant effect of married status disappears when the model is estimated free of selectivity bias, while the remaining coefficients continue to be largely unaffected. Thus, despite the apparent lack of significance of the selectivity term itself, these estimates indicate that marital status per se . . . fails to significantly shift the earnings profile" (p. 262).

This conclusion seems unjustified. The only real change in going from the single equation model to the model with endogeneity is a large increase in the standard error of the marital status coefficient. This seems an insufficient reason to discard the point estimate of the effect of marital status, which is, after all, a consistent estimate. In addition, their results indicate that the statistical experiment may be flawed.[6]

In summary, studies generally find substantial and statistically significant marriage wage premiums for males. The authors differ most widely in the interpretation of their empirical findings. The review suggests to us that, despite the contributions that these studies have made to a better understanding of marital pay differentials, continued research is needed to evaluate and refine hypotheses that are consistent with marital pay differentials. The remainder of the paper is devoted to this task.

III. NLS Data and Summary Statistics

The National Longitudinal Survey (NLS) of Young Men (Center for Human Resources Research, 1984) followed men aged 14 to 24 in 1966 for fifteen years. Data from 1976, 1978, and 1980 are used

6. A potentially important problem is that the selectivity term is itself correlated with marital status, by construction. This source of collinearity may lie behind the large increase in the standard error of the marital status coefficient, as well as the imprecision with which the coefficient of the selectivity term is estimated. Suggestive of other problems are: very large returns to experience (25 percent per year), negative returns to schooling, and the use of annual earnings with a sample of 16–24 year olds, many of whom may not have completed their schooling.

to obtain repeated observations on individuals,[7] which are required for estimation of the fixed-effects models presented below. We use the later years to capture the post-schooling labor market experience of a large portion of the sample. (The youngest men were 24 years old in 1976.) The sample is restricted to white men who completed schooling by 1976, and for whom all needed variables are available for the three years.[8] Attrition between 1966 and 1980 reduced the original sample of 5,225 men to 3,438. Data requirements further reduced the sample to 1,541; of these 1,228 are white.[9]

Sample statistics appear in Table 1. The first row of the table presents mean hourly wages in dollars according to marital status (the dependent variable in the regressions that follow is the natural logarithm of the hourly wage). Wages are available as hourly wages for hourly workers, and are constructed from weekly or annual earnings divided by the appropriate hours measure for those who report weekly or annual earnings. Never-married men appear to have much lower wages than men in the two other marital status groups (married, spouse present; and divorced or separated).

The figures in Table 1 also indicate that non-wage characteristics differ according to a man's marital status. For example, single men in the sample are younger and have completed more years of schooling, whereas married men have accumulated more work experience.[10] On average, married men worked about two hours more per week and one week more per year than never married men in the sample. The sample captures a fair number of marital status changes. Nearly one-third of the men who were single in 1976 had married by 1980 (29.1 plus 2.1, using figures for 1980 marital status from the third column), over one-half of the men who were divorced or separated in 1976 were "remarried" by 1980, and at least 8 percent of married men had separated from their wives by 1980.

7. Surveys were not conducted in 1977 and 1979.

8. Griliches (1976, 1978) discusses problems in using these data for younger men. While our choice of 1976–80 permits us to study post-schooling labor market behavior, it entails some reduction in the sample size and the number of marital status changers.

9. As mentioned in the introduction, Korenman and Neumark (1988) report results for blacks. We limit the present analysis to white men because the company personnel data set analyzed below contains no information on blacks. Separate analysis by race also seems advisable given evidence that patterns of family formation and marriage differ greatly between black and white Americans (Bennett, Bloom, and Craig 1989; Evans 1986).

10. "Actual Experience" measures post-schooling work experience, in years or fractions of years, based on weeks worked. This measure includes work experience of those who were enrolled in school part-time, but who did not report school as their primary activity. It was constructed using retrospective job history questions, and where these were incomplete, contemporaneous survey questions, whenever possible.

Table 1
Summary Statistics for White Young Men Classified by Marital Status in 1976

	Marital Status in 1976		
Characteristics in 1976	Married Spouse Present	Divorced or Separated	Never Married
Hourly wage ($)	6.57	6.90	5.56
	(0.08)	(0.40)	(0.17)
Schooling completed (years)	13.6	13.6	14.7
	(0.1)	(0.3)	(0.1)
Actual work experience (years)	6.6	6.2	4.2
	(0.2)	(0.6)	(0.3)
Age (years)	28.8	30.0	27.1
	(0.1)	(0.4)	(0.2)
Lives in South (%)	30.7	25.3	32.3
Lives in urban area (SMSA) (%)	70.6	84.8	79.9
Is covered by collective bargaining (%)	33.8	40.5	23.2
Has nonspouse dependent(s) (%)	77.6	51.9	3.7
Usual hours worked per week	43.7	43.6	41.5
	(0.2)	(1.1)	(0.6)
Weeks worked per year	50.7	48.8	49.5
	(0.2)	(1.0)	(0.6)
Years married	7.7	5.0	NA
	(0.1)	(0.4)	
Years divorced or separated	0.2	3.4	NA
	(0.0)	(0.3)	
Marital status in 1978, 1980 (%)			
Married, spouse present 1978	94.6	35.4	21.1
Married, spouse present 1980	92.0	53.2	29.1
Divorced or separated 1978	5.4	64.6	0.5
Divorced or separated 1980	7.9	46.6	2.1
Sample size	960	79	189

Data Source: National Longitudinal Survey, Young Men's Cohort.

The figures in Table 1 also serve as a reminder that divorced and separated men have spent a substantial number of years married.[11]

IV. NLS Empirical Findings

A. Methods

A leading explanation of marriage wage differentials is the selection hypothesis, which suggests that omitted, unmeasured characteristics that lead to higher wages are also positively correlated with marriage. For expository purposes, define the "true" model as

$$(1)\quad \ln(W_{it}) = \alpha \cdot X_{it} + \gamma \cdot MST_{it} + A_i + \epsilon_{it}$$

where W_{it} is the wage of individual i in year t, A_i is an unobserved characteristic of individual i, which is assumed to be time-invariant (for illustrative purposes only one marital status variable (MST), and one other observable characteristic, (X), are shown).[12] The selection of men with wage-enhancing attributes into marriage suggests that $\mathrm{Corr}(MST_{it}, A_i) > 0$. In this case $\hat{\gamma}$, estimated by least squares from a model in which A_i is part of the error structure, is an upwardly biased estimate of γ. A standard solution to this problem (e.g., Freeman 1984; Nakamura and Nakamura 1985) is "within" or fixed-effects estimation of the model:

$$(2)\quad \ln(W_{it}) - \ln(W_{i\cdot}) = \alpha \cdot (X_{it} - X_{i\cdot}) + \gamma \cdot (MST_{it} - MST_{i\cdot}) + \nu_{it}$$

where, for any variable Z, $Z_{i\cdot}$ denotes the mean of Z for individual i across the years t of the survey.[13]

Although estimation of this model removes individual fixed-effects, the possibility remains of serially correlated errors for an individual. Estimation of fixed-effects models in the presence of serially correlated disturbances is not trivial. Since in most panel data sets the number of periods is small, a consistent estimate of the residual covariance matrix is unavailable. We therefore use a consistent GLS estimator proposed by Keifer

11. Construction of the "years married" and "years divorced or separated" variables entailed a combination of contemporaneous and retrospective questions. For men who married for the first time prior to the first survey year, retrospective questions on the number and durations of marriages, divorces, and separations, allowed construction of marital histories covering up to two marriages prior to 1966. Fewer than 20 percent had been married at least once by the 1966 survey.

12. In the empirical analysis, the model will be expanded to include many "human capital" variables, as well controls for different marital states, and in some specifications, for the number of years spent in each marital state.

13. We often refer to the results from such models as longitudinal estimates.

(1980) that yields consistent estimates of coefficients and their standard errors. This estimator assumes an error structure where residuals are uncorrelated across individuals, but correlated across time for each individual. More specifically, stacking the data so that observations for each individual are grouped together, the covariance matrix of the residuals is $(I_N \otimes \Sigma_T)$, where N is the number of individuals, T is the number of periods, I_N is the $N \times N$ identity matrix, Σ_T is a symmetric, positive semi-definite $T \times T$ matrix, and $\otimes$ is the Kronecker product operator. Forming the "within" differences as in Equation (2) makes Σ_T a singular matrix, so that the GLS estimator requires a generalized inverse. Keifer shows that such an estimator can be obtained by forming the within differences, dropping data for one year,[14] and computing the GLS estimates, using the covariance matrix $(I_N \otimes \Sigma_{T-1})$.

B. Empirical Findings

Estimation of the potentially misspecified (pooled) cross-sectional wage Equation (1) (with A_i "omitted") confirms the presence of marriage premiums in the data. The first column of Table 2 presents these cross-sectional estimates, which closely resemble estimates from earlier studies. Married white men with spouse present earn about 11 percent more per hour than never-married men, controlling for survey year, labor market experience and its square, completed schooling (five dummy variables), union (covered) status, year of birth (ten dummy variables), South and urban residence, the presence of non-spouse dependents, and single-digit occupation and industry (nineteen dummy variables). Divorced or separated men earn about 9 percent more than never-married men.[15] Coefficients of these two marital status dummy variables are statistically significant at conventional levels.

Regressions (not reported) that also included years of service on the current job, and hours and weeks worked, yielded nearly identical results. Moreover (also not shown), dropping the control for nonspouse dependents raises the married coefficient somewhat, but leaves the divorced coefficient unchanged; and dropping the occupation and industry controls has no noticeable effect on the marital status coefficients.[16] This latter finding provides partial evidence against the hypothesis that marital status simply reflects unmeasured worker or job characteristics, since

14. We drop data for 1976, although the results do not depend on the year chosen.
15. The category divorced or separated includes a few men who are widowed or "married, spouse absent."
16. These additional regression results are available upon request from the authors (or see Korenman and Neumark 1988).

Table 2

Estimates of Marriage Premiums from NLS Wage Regressions, White Males, 1976, 1978, 1980. Dependent Variable Equals Ln(Hourly Wage) or Change in Ln(Hourly Wage); GLS Coefficients (standard errors in parentheses)

	Dummy Variable Specifications		Years Married Specifications	
	Cross-sectional[a] (1)	Longitudinal[b] (1′)	Cross-sectional[a] (2)	Longitudinal[b] (2′)
Married, spouse present	0.11	0.06	0.04	0.03
	(0.02)	(0.03)	(0.03)	(0.03)
Divorced or separated[c]	0.09	0.04	0.05	0.04
	(0.03)	(0.04)	(0.03)	(0.04)
Years married, spouse present			0.023	0.022
			(0.006)	(0.009)
Years married, squared/100			−0.085	−0.096
			(0.028)	(0.034)
Years divorced or separated			−0.011	−0.025
			(0.006)	(0.011)
Has nonspouse dependents	0.04	0.02	0.02	0.00
	(0.02)	(0.02)	(0.02)	(0.02)
Degrees of freedom	3,639	2,427	3,636	2,424

a. Also included in the cross-sectional regressions are actual experience and its square, South, urban, union (covered), schooling (5), year (3), year of birth (10), and single-digit occupation (8) and industry (11) dummy variables.

b. Except for schooling and cohort dummy variables, the longitudinal regressions include the controls listed above, entered as deviations from their mean values from 1976, 1978, and 1980. Because the sample is restricted to those who have completed school, schooling variables do not deviate from their mean values and are excluded. Birth year is also constant over time.

c. The category divorced or separated includes a few men who are widowed or married, spouse absent.

the occupation and industry dummy variables should control for average differences in these characteristics across (broad) occupations and industries.

Column (1′) of Table 2 presents coefficients from longitudinal (fixed-effects) estimation of Equation (2). Men who marry experience about an 8 percent wage increase, while those who divorce experience roughly a 2 percent decrease (e.g., 0.04 minus 0.06). Clearly the longitudinal estimates (and the implied t-statistics) of the effects of marriage are much smaller than the cross-sectional estimates. The much lower estimates of the marriage premium found in the longitudinal analysis are consistent with the hypothesis that a good portion (roughly one-half) of the cross-sectional marriage premium is associated with fixed (econometrically) unobservable characteristics of individuals that are positively correlated with both marriage and wages. The differences between the cross-sectional and longitudinal marriage premiums, however, could also result from an important misspecification of the earnings equation that could contaminate the statistical experiment. In particular, it is possible that the attenuation of the marriage premium results from a time-dependent effect of marriage on wages (as suggested by Kenny 1983). In contrast to the benefits of starting a unionized job, for example, that should accrue to a member quickly, there is no reason to believe that the labor market returns to marriage are reaped upon utterance of the words "I do." This consideration suggests that the diminution of the marriage premium in longitudinal estimation may be due to a shorter duration of marriage for those who marry between 1976 and 1980 than for those who are observed married at a given time. In our sample, a person who married between 1976 and 1978 accumulated about one and one-half years of marriage, on average, while those who were married in 1978, for example, had been married for an average of about twelve years. Therefore, a longitudinal marriage wage differential of 6 percent is consistent with a 4 percent yearly wage growth premium (in the early years of marriage) accruing to married, as compared to never-married men.

Columns (2) and (2′) of Table 2 report coefficients from cross-sectional and longitudinal regressions where marriage premiums are allowed to vary by marital duration.[17] The most striking result is that there is no longer a sizable or significant *intercept* shift associated with any particular marital status. Rather, the effect of marriage appears gradually, increasing wages (cross-sectionally) 2.3 percent per year married in the early years of marriage, and roughly 1–2 percent per year at the mean of years

17. Corresponding sample statistics by years married are presented in Table A1 in the appendix.

married.[18,19] The wage premium earned by divorced or separated men compared to never married men appears to be explained by advantages gained from time spent married; there is a negative and significant cross-sectional effect on wages of years divorced or separated. The fixed-effects results also reveal a significantly negative association between years divorced and wages of 2.5 percent per year (compared to continuously never-married men).

Calculations of marital premiums for persons who change marital states lend additional support for the idea that the reduction of the marital status premium from 0.11 in Column (1), to 0.06 in Column (1′) is due to a misspecification of the wage equation. As noted, a man who married between 1976 and 1978 accumulated 1.5 years of marriage, on average. Using this figure in conjunction with the coefficients presented in Column (2) yields a marriage premium for these newlyweds of 0.07, and the coefficients reported in Column (2′) imply a premium of 0.06,[20] very similar to the fixed-effects coefficient reported in Column (1′). Columns (2) and (2′), however, suggest that the marriage premium continues to grow with each year married, indicating that the 0.06 percent figure understates the marriage wage differential. This point is brought out more fully by Table 3.

In Table 3, we use the cross-sectional and longitudinal coefficient estimates from Table 2, Columns (2) and (2′), to summarize the marriage wage premium in two ways: first, as the sum of the coefficient on the dummy variable plus the years effect (evaluated at the unconditional sample means for the appropriate group); and second, as this figure, plus the coefficient on the dependents dummy. The latter calculation compares a married man with nonspouse dependents to a single man with no dependents. (For comparison, corresponding figures from the dummy variable specifications in Columns (1) and (1′) of Table 2 are reproduced in the second panel of Table 3.) Table 3 indicates that the longitudinal estimates are much closer to the cross-sectional estimates for the "years married" specification than for the dummy variable specification. Evaluated at the sample mean years of marriage, the longitudinal estimate of the marriage premium is between 80 and 90 percent as large as the cross-sectional estimate.

18. The quadratic functional form for years married was adopted after examining specifications with single year dummy variables for each year of marriage.

19. In regressions that included both the total duration in each marital state and the duration in the current marital state, duration in current state did not enter significantly.

20. Using the coefficients reported in Column (2′), the relevant calculation is

$$0.06 = 0.03 + (1.5) \cdot (0.022) + (1.5)^2 \cdot (-0.096)/100.$$

Table 3

Summary of Cross-Sectional and Longitudinal Estimates of Marital Pay Premiums at Mean Years Married, NLS Young Men, Whites, 1976, 1978, 1980. Estimated Wage Premiums (Percents)[a]

	Married versus Single			Married with Nonspouse Dependents versus Single with No Nonspouse Dependents		
	Cross-Section	Longitudinal	Ratio	Cross-Section	Longitudinal	Ratio
Years married specification[b]						
Table 2, column (2) versus (2′)	16.6	14.8	.92	18.1	15.2	.84
Dummy variable specification						
Table 2, column (1) versus (1′)	10.6	6.1	.58	14.8	8.2	.55

a. Estimates of wage premiums are calculated using coefficient estimates from Table 2. For regression details, see footnotes to Table 2.
b. Estimates are evaluated at the unconditional sample mean of "Years Married" for the appropriate group.

To summarize findings from the NLS data, there appears to be a significant increase in wages associated with marriage for white males, even correcting for selectivity into marriage based on fixed unobservables; over eighty percent of the estimated impact of marriage on earnings survives the fixed-effects estimation. Moreover, large marriage premiums persist even after adding controls to wage equations that should capture differences across marital statuses in labor supply or in investment in human capital (such as occupation, industry, hours and weeks worked, and tenure on the current job).

Although the fixed-effects estimates suggest that the selectivity hypothesis alone does not account for a major share of the marriage premium, at least not in the simple "fixed unobservables" form of the hypothesis, it is possible that a more complex selection process underlies the marriage premium. For example, selection into marriage on the basis of unmeasured potential wage growth could account for the remaining marriage premiums. The possibility also exists that causation could run from wages to marital status. If the marriage premium resulted from selection of men with high wage *levels* into marriage, then we would not expect to have found that marriage wage premiums result from wages rising with years married. On the other hand, if the marriage wage growth premiums were due to selection of men with high wage *growth rates* into marriage, then unmarried men with relatively high wage growth in a period should be more likely to marry thereafter. Korenman (1988, Chapter III) found no evidence of such an effect in these data; controlling for other observable characteristics, unmarried men in this sample who had high wage growth between 1975 and 1976 were (insignificantly) *less* likely than other men to marry between 1976 and 1980.[21]

In the remainder of the paper we present analyses of data drawn from a company personnel file in order to illuminate further the marriage differential in general, and the "years married" effect in particular.

V. Company Personnel Data and Empirical Findings

A. Data

This section briefly describes data from the personnel file of a large U.S. manufacturing firm described in detail in Medoff and Abraham (1981). The data are for white male managers and professionals working within a single firm for the year 1976. Sample statistics are reported in Appendix

21. Tables summarizing these estimates are available upon request from the authors.

Table A2. Because the data pertain to a fairly homogeneous set of occupations (managers and professionals) and work environments (a single firm), they allow us to control implicitly for important characteristics of workers and jobs that potentially vary across persons of different marital status. Most of the variables are self-explanatory as they are common in labor market data sets.

The data are unique in that they obtain supervisor performance ratings that provide an additional measure of worker productivity, aside from a worker's wage. The performance of each manager or professional employee is rated each year by his immediate supervisor.[22] Supervisors are instructed to rate employees on "current performance and contributions based on requirements of his present assignment" (Medoff and Abraham 1981, p. 189). Ratings are to reflect performance relative to the "standards" of the job, and relative to "others performing similar work at similar levels," but are not to reflect "[c]areer potential and promotability" (Medoff and Abraham 1981, p. 189). On a six point scale, six is the highest performance rating. Each employee is also ranked relative to other workers in comparable jobs, by several different supervisors, all of whom are familiar with his work. In order to ensure the consistency of performance ratings by immediate supervisors with these "consensus rankings," both are reviewed by managers at the next higher level of the corporation.

Another somewhat nonstandard variable contained in the data is the employee's job grade. According to a company salary manual, a job's grade reflects its relative value to the company, and "positions of similar value are placed in the same classification level" (Medoff and Abraham 1981, p. 191). (In practice, job grade dummy variables alone account for roughly 85 percent of the variation across workers in the log of annual salaries in this company.)

Given these definitions of performance ratings and job grades, Medoff and Abraham are "comfortable assuming that, within a given grade level, those with high performance ratings or high rankings were more productive than those with low performance ratings or rankings" (Medoff and Abraham 1981, p. 191).

B. Estimation and Empirical Findings

Table 4 presents estimates from earnings equations in 1976, using the natural log of annual salary as the dependent variable. The second column shows sizable and significant annual earnings differentials of about 12 percent favoring both married and divorced men, controlling for pre-

22. This section summarizes, in part, Section I of Medoff and Abraham (1981).

Table 4
Estimated Coefficients from Cross-Sectional Earnings Regressions. Company Personnel Data, White Male Managers, and Professionals.[a,b]
Dependent Variable Equals Ln(Annual Salary as of December 1976)

	OLS Coefficients (SEs)			
	(1)	(2)	(3)	(4)
Married	.226	.119	.025	.019
	(.014)	(.012)	(.004)	(.004)
Divorced	.182	.115	.022	.022
	(.027)	(.021)	(.008)	(.007)
Widowed	.224	.069	.020	.013
	(.042)	(.034)	(.013)	(.012)
Other Included Variables				
Precompany experience and its square,				
Company service and its square,	no	yes	yes	yes
Region (3) and education (4) dummies				
Grade dummies (11)	no	no	yes	yes
Performance rating dummies (5)	no	no	no	yes
R^2	.03	.38	.91	.93

a. The sample size is 8,235. The reference category is "Single."
b. Means of variables used in the regressions appear in Table A2.

company experience and its square, company service and its square, and dummy variables for region and education. Because performance ratings are valid only for workers doing comparable jobs (i.e., controlling for job grade), performance ratings can be meaningfully related to worker productivity only if job grade dummies are added to the wage regressions. Adding job grade dummy variables alone (column three) diminishes the marriage premium substantially; although the coefficients remain statistically different from zero, they imply that married men earn only about 2.5 percent more per year than otherwise comparable single men working in similar jobs.[23] Thus most of the return to marriage, whatever its source, takes the form of location of married workers in higher paying job grades within the company, rather than higher pay for a given performance level within a job grade.[24]

The preceding results imply that we should direct our attention to the association between marriage and job grades—and not to the effects of marriage on wages within job grades. Why are single men relatively concentrated in the lower grades? On the one hand, marital status could determine the grades into which workers are hired or promoted, or whether they experience job separations. On the other hand, the concentration of married workers in higher grades may be coincidental, simply reflecting a company lifecourse whereby young workers are hired into lower grades, progress through the corporate hierarchy based on a combination of tenure and performance, and tend to get married along the way. If this last scenario is correct, then there should be no discernible marriage pattern of hiring, promotion, pay, or performance once controls are introduced for other worker characteristics and company tenure.

To explore the first possibility, we examine a sample of "recent hires" (workers with two or fewer years of company service). There is at most a slight marriage pattern to hiring; over 90 percent of recent hires are found in four job grades (2–5); single workers comprise 32 percent of the recent hires in grade two, and 29 percent of recent hires in each of grades three through five. Further, estimates of earnings equations for this sample, presented in Table 5, confirm that hiring practices are not responsible for the marriage premiums. In particular, there is no significant wage differential favoring married "recent hires," controlling for the worker characteristics listed above, whether or not grade dummy variables or performance ratings are included in the regressions.[25]

Turning to the possibility of different chances of promotion into higher grades, or job separation, on the basis of marital status, we estimate multinomial logit models of the probability of promotion and job separation of recent hires between year-end 1976 and year-end 1977, including

23. These regressions were also run separately for the 12 individual job grades. The coefficient on the married dummy variable changed somewhat from grade to grade, but was generally about the same size as the coefficient for the overall sample, controlling for job grade. There was no striking correlation between job grade and the size of the marriage premium.

24. That single workers are disproportionately represented in lower job grades is apparent from the job grade distribution reported in Table A2.

25. The sample was restricted to four grades that contain over 92 percent of the company's "recent hires," as well as to men who were married or single (eliminating two divorced men), and who received better than the lowest performance rating (eliminating one employee).

Table 5
Estimated Coefficients from Cross-Sectional Earnings Regressions. Company Personnel Data, White Male Managers and Professionals, Recent Hires.[a,b,c] *Dependent Variable Equals Ln(Annual Salary as of December 1976)*

	OLS Coefficients (SEs)			
	(1)	(2)	(3)	(4)
Married	.030 (.015)	−.004 (.013)	.002 (.010)	.000 (.010)
Other included variables				
Precompany experience and its square				
Company service and its square	no	yes	yes	yes
Region (3) and education (3) dummies				
Grade dummies (3)	no	no	yes	yes
Performance rating dummies (4)	no	no	no	yes
R^2	.01	.40	.63	.64

a. N = 280. The reference category is "Single."
b. "Recent Hires" are people with two or fewer years of company service.
c. The smaller number of grades here than in Table 3 reflects the concentration of "Recent Hires" in the lower grades. Only one recent hire received the lowest performance rating, hence he and the category were dropped from the sample.

controls for characteristics of workers as of 1976. Results are summarized in Table 6. In the first panel, where performance ratings variables are excluded from the analysis, marriage is associated with a substantially higher probability of promotion, and a slightly lower probability of separation from the company, controlling for precompany experience and its square, company service, and dummy variables for education, region, and job grade. Although neither coefficient is statistically significant (the asymptotic t-statistic for the promotion-marriage coefficient is about 1.5, and for the job separation-marriage coefficient, about 0.6), the implied partial derivative of the probability of promotion with respect to marriage is sizable; controlling for the above characteristics, the probability of promotion for a married man is 10.5 percentage points higher than for a single man (Column 2). Compared to the sample mean probability of

Table 6
Estimated Coefficients from Multinomial Logit Models of Promotion and Job Separation, and Models of Rated Performance, Recent Hires. Company Personnel Data, White Male Managers and Professionals, 1976 (standard errors in parentheses).[a]

	Married Coefficient (SE)	Partial With Respect to Marriage (percent)	Percent Change in Probability (at sample mean)
1. Without performance controls[b,c]			
Job separation	−.223 (.378)	−2.8	−16.9
Promotion	.499 (.332)	10.5	34.3
2. With performance controls			
Job separation	−.318 (.339)	−4.1	−24.7
Promotion	.052 (.368)	1.0	3.4
3. Performance rating[d]			
Low	−.636 (.358)	−13.2	−43.0
High	.791 (.355)	19.3	46.2

a. "Recent Hires" have two or fewer years of company service. The sample size is 280. The partial derivatives are evaluated at sample means.
b. Other variables included in the promotion and separation equations are: precompany experience and its square, company service, and dummy variables for education (3), region (3), job-grade (3), and a constant. Performance controls are four rating dummy variables.
c. In the sample, 16.4 percent of recent hires had job separations, 30.7 percent were promoted. The reference category is "retained without promotion."
d. Other variables included in the performance equations are: precompany experience and its square, company service, dummy variables for education (3), South, job-grade (3), and a constant. In the sample, 30.7 percent were rated "Low" and 41.9 percent were rated "High," i.e., below or above the median (reference) category.

promotion, married men are about 34 percent more likely to be promoted (41.2 versus 30.7), all else (except rated performance) held constant.[26]

Coefficients from analyses that include performance dummy variables are reported in the second panel. Strikingly, the marriage promotion premium virtually disappears when performance ratings are added; the separation premium also grows somewhat (in absolute value). As these figures suggest, and as multinomial logit analyses of performance ratings confirm (Panel 3), married workers are more likely to win high performance ratings, and higher performance ratings are positively related to the probability of promotion.[27]

The third panel of Table 6 summarizes results from these performance rating logits (the top two and bottom two performance rating categories are combined). The figures indicate that marriage increases by almost 50 percent the probability that a recent hire will receive one of the top two performance ratings (compared to the median rating category), controlling for education, job grade, Southern location, company service, and precompany experience.

In short, married men are substantially more likely than their single counterparts to receive high performance ratings; high performance ratings, in turn, appear to increase promotion probabilities so that married men are also more likely to be promoted. The large and statistically significant marital pay premiums for managerial and professional workers within a single firm suggest that marriage wage differentials do not compensate married men for working under adverse conditions. While it is possible that these managers and professionals need to be compensated for undertaking the greater responsibility that accompanies promotion, it seems plausible that promotions to higher paying positions are used to reward workers who are perceived to be performing well.

These results also provide evidence that is consistent with the NLS data, which indicate that a steeper wage profile, rather than an intercept shift, best describes marriage wage premiums. The evidence presented therefore calls for future research to focus on job assignment or promotion, in additon to wages, in exploring marriage pay differentials. Because these findings are based on analysis of the company personnel file of a

26. This finding is consistent with Goldin's (1990, p. 102) evidence from the Depression era suggesting that employers promoted married office workers more frequently.

27. In particular, the promotion logits indicate large and statistically significant relationships between the probability of promotion and the four included performance rating dummy variables (2, 3, 4, and 6, there are no "1s" in the sample, 5 is the reference category). The derivatives of the probability of promotion with respect to these four dummy variables are: −21.4, −45.0, −16.5, and 16.7; the ratios of the coefficients to standard errors (corresponding to the coefficients used to calculate the derivatives) are: −1.5, −4.4, −2.1, and 1.7, respectively.

single firm, caution should be exercised in drawing more general conclusions from these data until additional firm-level analyses are conducted.

VI. Concluding Remarks

In this paper we have attempted to bring new evidence to bear on the male marital pay premium. We have five principal findings to report.

First, hourly wage premiums paid to married men are large and persist even when detailed human capital controls (such as actual labor market experience) are included in wage equations.

Second, marriage premiums seem to arise slowly, resulting more from faster wage growth for married men as compared to never-married men, than from an intercept shift associated with any particular marital status. This finding suggests that an appropriate specification allows for a relationship between marriage and wages that varies according to the number of years a man is married.

Third, taking as our "true" model a specification that incorporates a "years married" effect, comparisons of cross-sectional and fixed-effects estimates indicate that selection on the basis of fixed unobservable characteristics accounts for less than twenty percent of the observed wage premium.

Fourth, data drawn from a company personnel file indicate that the marriage wage premiums persist even within a single firm, for a relatively homogeneous group of occupations (managers and professionals). The marriage premium appears to be due to the location of married workers in higher paying job grades within the company, rather than to married workers receiving higher pay than single workers within the same job grades.

Finally, married workers in this company receive higher performance ratings from their supervisors. These higher performance ratings increase their promotion chances, allowing them to enter the higher paying job grades. When we controlled for supervisor performance ratings, however, the marriage promotion premium disappeared.

Although these findings do not represent a tightly constructed test of a theory of marital pay differentials, they do lend greater support to some hypotheses than to others. In particular, selection of men into marriage on the basis of wages, wage growth, or other wage-enhancing characteristics receives little support as an explanation of the observed marital pay premiums. Wages rise after marriage, fixed-effects estimates of marriage premiums in properly specified wage equations are nearly as large as their cross-sectional counterparts, and in the sample and period we studied, unmarried men's wage growth appears unrelated to the probability that they marry.

While it may be fair to interpret our findings as providing support for the hypothesis that marriage enhances men's labor market productivity, a definitive judgment requires more direct evidence about the processes or causal mechanisms linking marriage to men's productivity.

Table A1
Sample Statistics by Years Married, Ever Married Men, 1976, Means (standard errors of means in parentheses).

Characteristics in 1976	Two or Fewer	Two to Five[a]	Five to 10	10 or More
Hourly wage	5.51	6.02	6.77	7.22
	(0.36)	(0.13)	(0.13)	(0.17)
Schooling completed (years)	13.9	14.1	13.8	12.7
	(0.3)	(0.1)	(0.1)	(0.1)
Actual work experience (years)	3.9	4.4	6.3	10.3
	(0.5)	(0.2)	(0.2)	(0.3)
Age (years)	26.9	26.5	28.7	32.1
	(0.3)	(0.1)	(0.1)	(0.1)
Percentage who				
Live in South	27.1	25.0	33.6	31.5
Live in urban areas (SMSAs)	83.1	74.7	69.1	70.1
Are covered by collective bargaining	47.5	29.8	35.3	34.6
Have nonspouse dependents	35.6	47.6	86.6	98.4
Hours worked per week	42.5	42.8	44.4	43.8
	(1.1)	(0.4)	(0.4)	(0.4)
Weeks worked per year	49.2	50.0	50.8	51.0
	(1.3)	(0.4)	(0.3)	(0.3)
Married, spouse present (%)	64.4	93.2	92.4	98.0
Divorced or separated (%)	35.6	6.8	7.6	2.0
Years married spouse present	1.1	3.7	7.7	13.0
	(0.1)	(0.1)	(0.1)	(0.1)
Years divorced or separated	1.8	0.4	0.4	0.2
	(0.4)	(0.1)	(0.1)	(0.1)
Sample size	59	292	434	254

a. Includes 5.0 years

Table A2
Means for White Male Managers and Professionals, Company Personnel Data, 1976 (standard deviations in parentheses).

	All	Single	Married	Divorced	Widowed
Number	8,235	337	7,730	126	42
Salary in 1976 ($)	26,629	21,384	26,873	25,594	26,936
	(7,183)	(5,731)	(7,159)	(6,347)	(8,119)
Ln(annual salary)	10.15	9.94	10.17	10.12	10.16
	(0.25)	(0.24)	(0.25)	(0.23)	(0.27)
Less than high school diploma	.02	.01	.02	.01	.10
High school diploma	.34	.09	.35	.36	.43
College	.45	.52	.44	.42	.38
MA	.15	.29	.14	.18	.10
PhD	.04	.09	.04	.03	.00
Years of precompany experience	5.2	2.9	5.3	4.9	5.9
	(5.3)	(3.0)	(5.4)	(4.9)	(5.9)
Years of company service	19.5	8.5	19.9	16.4	31.1
	(11.3)	(10.4)	(11.0)	(10.9)	(8.2)
Northeast	.14	.14	.14	.06	.29
South	.75	.71	.75	.81	.67
Midwest	.02	.03	.02	.02	.00
West	.09	.13	.09	.10	.05
Job grade distribution					
Bottom quartile	11.5	17.2	11.3	11.9	19.1
Second quartile	51.5	65.9	50.8	55.5	47.7
Third quartile	30.2	14.9	31.0	26.9	23.8
Top quartile	6.8	2.0	6.9	5.6	9.6
Performance Ratings					
Low (1–3)	40.1	38.0	40.2	39.7	52.4
High (4–6)	59.9	62.0	59.8	60.3	47.6

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