9-27-1992

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DIFFERENCES IN LONGITUDINAL UNION RELATIVE WAGE EFFECTS ACROSS GENDER AND RACE

by

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Economic Staff Paper No. 92-10**

September 1992

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Sixty-six copies of this document were printed by ghe Economics Department at a cost of $1.26 per document.

ABSTRACT

This paper utilizes longitudinal union variables to examines the gender differences in the wage change from entering and leaving the union sector, as well as gender differences from remaining with a union employer and advancing up the seniority ladder. The union effects are estimated over a time trend from the late 1960s to the early 1980s. The empirical results show that the female union joiner effect declines over time while the magnitude of the male union effect remains fairly stable over time. In general, the results show that unions' ability to impact wages for all white workers and black female union joiner has attenuated over time while union effects for all other black workers have remained relatively constant.

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INTRODUCTION

Due to a persistent gender gap of 30% to 40%, women face a lower wage structure and have an incentive to seek union sector employment which typically pays higher wages. To the degree that the gender wage gap results from labor market discrimination, unions may help ameliorate women's pay differential since collective bargaining agreements have formal rules for pay, on-the-job training, and promotion. Researchers, over the years, have attempted to ascertain whether collective bargaining benefits female workers more than male workers. That is, is the union wage gain larger for females or males, relative to their nonunion counterparts? Lewis [21] reviews many of the studies on gender variations in the union wage differential and finds that the union differential is the same, on average, for males and females. Black males, on the other hand, generally receive a larger union differential than white males. Consequently, unions have been criticized as a hinderance to the improvement of the female wage distribution. Much of the criticism has been directed at seniority ladders. Since union jobs, particularly in high wage industries, are dominated by males, the seniority ladders prevent women from advancing into high wage occupations, and cause them to incur proportionally greater lay-offs during economic downturns. However, as seniority increases, women's advancement through occupational job ladders should enhance their wage distribution. Freeman and Leonard [13] find larger female differentials in the public sector and in some industries dominated by females. Figart [10] claims that unions have increased women's wages more than men's over the 1980s, especially in the 35 to 54 age bracket.

Traditionally, union status is measured by a single period dichotomous union variable. The parameter estimate for the union dummy variable is interpreted as the average relative wage differential from collective bargaining, and it does not differentiate between workers who have recently entered the union sector and workers who have remained in the union sector to advance up the seniority ladders. The interpretation of the cross-sectional union wage gap as resulting solely from collective bargaining has been criticized by numerous authors (e.g., Duncan and Leigh [9], and Lee [18]). Unmeasured or unobservable differences in labor quality may cause some or all of the wage gap since union employers would have an incentive to offset increased labor costs by hiring the most productive workers from the labor pool. Therefore, it is likely that the union wage gap is positively correlated with unobserved productivity differences. The pioneering studies by Mellow [22] and Mincer [23] overcome the problems associated with unobserved quality characteristics by estimating the worker-specific wage change from entering (or leaving) the union sector over a longitudinal period. The longitudinal union effect will not be biased by unobservable productivity characteristics since it is based on an individual's wage in both the union and nonunion sectors. The longitudinal model developed by Mincer adds mobility among employers to the variables measuring change in union status. Mincer's model provides a more accurate assessment of variation in the union effects by gender, since it isolates workers searching for new jobs from those who remain with their initial employer.

This study utilizes the Mincer longitudinal union status variables to investigate the gender differences of entering and leaving the union sector, as well as the gender differences in the wage change from remaining with a union employer and advancing up the seniority ladder. Previous studies rely on the traditional union variable to estimate union effects and report results for only one or two time periods. It is conceivable that the union differential changes over time. This study reports a time trend over 1970s and early 1980s to illustrate the gender variation in union effects. The data set, the Mincer wage level and wage change models, and the decomposition of the wage change into the "true" union entry effect are discussed in the second section. The third section discusses the empirical results and compares the time trend of white and black female union effects with white and black males, and the study is concluded in the fourth section.

DATA AND METHODOLOGY

The data source is the National Longitudinal Survey (NLS) of young women and young men. The NLS young women and men data sets were selected because younger workers are more mobile between jobs and therefore have a higher probability of changing their union status. The data sets contain cohorts who were between the ages of 14 and 24 during the initial survey period of 1968 for young women and 1966 for young men.[1] Since the collective bargaining status is not reported in all of the years, this study utilizes the 1969, 1970, 1971, 1976, 1978, and 1980-81 young men's surveys; and the 1970-73, 1977-78, 1980, and 1982 young women's surveys. The young women's data are separated into seven longitudinal periods: 1970-71, 1971-72, 1972-73, 1973-77, 1977-78, 1978-80, and 1980-82. The young men's data are separated into six longitudinal periods: 1969-70, 1970-71, 1971-

76, 1976-78, 1978-80, and 1980-81. Separate wage equations are estimated for white females, black females, white males, and black males for each respective longitudinal period.[2] The longitudinal periods represented by this data span a variety of macroeconomic conditions and permit the estimation of wage differentials and wage change effects over different phases of the business cycle.[3]

Wage Level Model

The cross-sectional wage level equation used in this study is similar to Mincer [23], and has the following general form:

$$\begin{aligned} \ln w_{it} = \beta_0 + \delta_{01m_t} u_{01m_{it}} + \delta_{10m_t} u_{10m_{it}} + \delta_{11m_t} u_{11m_{it}} + \delta_{00m_t} u_{00m_{it}} + \\ \delta_{01s_t} u_{01s_{it}} + \delta_{10s_t} u_{10s_{it}} + \delta_{11s_t} u_{11s_{it}} + \sum_{k=1}^{K} \beta_k x_{itk} + \gamma \lambda_{it} + \mu_{it} \end{aligned} \tag{1}$$

where i indexes the individual cross-sectional observations, and t indexes the time period. The dependent variable is the natural logarithm of the real hourly wage rate. The x_{it}s represent the personal characteristics which are assumed to be nonstochastic. The human capital variables include: education, potential labor force experience, experience squared, seniority, and seniority squared. The other standardizing variables are: local area unemployment rate, marital status, health status, southern residence, residence in a metropolitan area, and longitudinal union status.[4] β_0 is the intercept, and the β_k are the slope coefficients. As the subscripts indicate, the intercept and the slope coefficients are assumed to be constant over individuals and time. μ_{it} is the error term where $\mu_{it} = \alpha_i + \epsilon_{it}$. α_i is the time invariant component of the error term which reflects unobservable individual-specific influences that help determine w_{it}; ϵ_{it} is the component of the error term that varies over individuals and time and is assumed to have a $N(0,\sigma^2)$ distribution. λ_{it} is the inverse of the Mills ratio from a probit equation that predicts the probability of sample inclusion. The inverse Mills ratio is added to the female regression to correct for sample selection bias.[5]

The Mincer model estimates the union impact on wages by including seven longitudinal union change variables. The data are first separated into two main groups: (1) those who move between employers between times t and $t+1$, (i.e., movers); and (2) those who remain with their original employer over the time period, (i.e., stayers). For the mover group, there are four union change variables: (1) u_{01m}, a union joiner; (2) u_{10m}, a union leaver; (3) u_{11m}, a union stayer; and (4) u_{00m}, a nonunion stayer. For the stayers group, there are also four union change variables: (1) u_{01s}, a union joiner; (2) u_{10s}, a union leaver; u_{11s}, a union stayer; and (4) u_{00s}, a nonunion stayer (the base group). Since this analysis focuses on the union effect of entry into and exit from union sector employers and the wage differential attributed to union job seniority, the study limits the empirical analysis to the u_{01m}, u_{10m}, and u_{11s} variables. The analysis includes the u_{00m} group since the results are needed to calculate the true union effect.

The interpretation of the coefficients on the union/nonunion change variables depends on whether equation (1) is estimated for period t or $t+1$. In period t when (1) is the *ex ante* wage equation, δ_{01m} estimates the nonunion wage differential between nonunion workers who will change employers to join the union sector in period $t+1$ and the nonunion base group, u_{00s}. In period $t+1$ when (1) becomes the *ex post* wage equation, δ_{01m} estimates the union wage differential relative to the nonunion base group. In period t, δ_{10m} estimates the union wage differential between workers who will leave the union sector and their current employer in period $t+1$ and the nonunion base group. δ_{10m} in period $t+1$ estimates the wage gap between union leavers and the base group, after both groups are in the nonunion sector. In period t, δ_{00m} estimates the wage gap between the base group and workers who will change employers but remain nonunion. δ_{00m} in period $t+1$ estimates the wage gap between nonunion workers' wages at their new employer and the base group. The difference between δ_{00m} in period $t+1$ and δ_{00m} in period t is the mobility premium from job search. In both periods, δ_{11s} estimates the union differential between the nonunion base group and workers who remain in the union sector with their original employer.

Since the union change variables measure union or nonunion status between two surveys, the variables will have the same value for the t and $t+1$ survey periods. The mean value for u_{11S} in Table 1 indicates that males have a greater union coverage rate than females.[6] The greater proportion of blacks in u_{01m} and u_{10m} indicates that they tend to be more mobile than whites in and out of the union sector. Females, however, have a greater presence in the nonunion sector as indicated by the higher proportion of females in u_{00m} and the base group.

Wage Change Model

The wage change model uses the first-difference of the $t+1$ and t cross-sectional variables for each respective longitudinal period, and reveals the wage change due to a change in union status from period t to $t+1$.

The wage change model has the following specification:

$$\Delta \ln w_{it} = d_{01m_i}u_{01m_{it}} + d_{10m_i}u_{10m_{it}} + d_{11m_i}u_{11m_{it}} + d_{00m_i}u_{00m_{it}} + d_{01s_i}u_{01s_{it}} + d_{10s_i}u_{10s_{it}} + d_{11s_i}u_{11s_{it}} + \sum_{k=1}^{K} \beta_k \Delta x_{ikt} + b\Delta\lambda_{it} + \Delta\epsilon_{it} \tag{2}$$

where it estimates the change in the natural logarithm of the real hourly wage as a function of changes in the same variables included in the wage level equation. The change in experience is equal for all individuals in each time period and becomes the intercept for the wage change model.[7] Since the wage change model is equivalent to the first-difference of the cross-sectional wage level equation, it eliminates α_i and the parameter estimates are free of heterogeneity bias. The d_{01m} coefficient estimates the wage change from entering the union sector after changing employers; d_{10m} estimates the wage change from leaving the original employer and the union sector and is interpreted as a negative union effect; d_{00m} estimates the wage change from changing nonunion employers and is interpreted as the mobility premium from job search; and d_{11s} estimates union negotiated wage changes.

The wage change model yields identical results to the first-difference of the $t+1$ and t period wage level equations. Mincer uses this relationship to determine the "true" union joiner effect which he argues is the wage change estimate of the union joiner effect, d_{01m}, minus the mobility premium from job search, d_{00m}. The true union joiner effect can be expressed by:

$$\begin{aligned}[(\delta_{01m} - \delta_{00m})_2] - [(\delta_{01m} - \delta_{00m})_1] &= [(\delta_{01m})_2 - (\delta_{01m})_1] - [(\delta_{00m})_2 - (\delta_{00m})_1] \\ &= (\Delta\delta_{01m} - \Delta\delta_{00m})\end{aligned} \tag{3}$$

where the subscripts *1* and *2* indicate that the parameter estimates are from the *ex ante* and *ex post* cross-sectional wage level equations, respectively. The first bracketed term on the left-hand side is the *ex post* new hire union wage difference between union joiners who changed employers and movers who stayed in the nonunion sector. The second bracketed term on the left-hand side is the selectivity bias term which is estimated from the *ex ante* period nonunion wage difference between union joiners and nonunion stayers. This term is interpreted as a productivity differential (based on wages in the current job) between prospective union workers and nonunion workers who also search for new jobs. The left-hand side of the equation indicates that the true cross-sectional union differential is net of selectivity bias.

After rearranging terms, the first bracketed term on the right-hand side becomes the first-difference estimate of the union joiner effect. Since heterogeneity bias is additive to the cross-sectional wage level coefficients, $(\delta_{01m})_2$ and $(\delta_{01m})_1$, the heterogeneity bias term is eliminated in the first-difference and the union joiner effect is unbiased. The second bracketed term on the right-hand side is the mobility premium. $\Delta\delta_{01m}$ is the wage change model's estimate of the union joiner coefficient, d_{01m}, and $\Delta\delta_{00m}$ is the estimate of the nonunion stayer coefficient, d_{00m}. The term $(\Delta\delta_{01m} - \Delta\delta_{00m})$ from the first-difference model corresponds to $d_{01m} - d_{00m}$ from the wage change model. These terms can be used to estimate the true return to union joiners, net of heterogeneity bias and the mobility premium.[8]

EMPIRICAL RESULTS

The female and male wage change parameter estimates are given in Tables 2 and 3, respectively; and the female and male wage level estimates are given in Tables 4 and 5, respectively.[9] In the wage change model, the coefficient values reflect the percentage wage change associated with the mobility pattern specified by a particular union status. Since the wage change model explains the percentage change in wage levels, the parameter estimates from the underlying *ex ante* and *ex post* wage level models reveal the source of the wage change phenomena.

Wage change parameter estimates for female u_{01m} workers (Table 2) suggest that both whites and blacks receive a positive, significant union joiner effect in the 1970-71 through 1973-77 time periods. In 1970-71, the white u_{01m} coefficient is .2198 (Table 2), implying that wages for white females increase by $e^{.2198}-1$ or 24.6 percent after joining the union sector. The black u_{01m} coefficient of .2259 has a similar interpretation. In Figure 1, both the white and black u_{01m} coefficients for the 1971-72, 1972-73, and 1973-77 periods have relatively similar magnitudes. By 1977-78, the female union effect begins to decline as the black effect is significantly smaller; and by 1980-82, neither female group receives a significant union joiner effect. The time trend in Figure 1 suggests that the female wage change from the union entry declines dramatically in the late 1970s and early 1980s.

The underlying cross-sectional wage level estimates for female union joiners imply the wage change from union entry. The u_{01m} wage level results (Table 4) suggest that the positive wage changes for the 1970-71, 1971-72, and 1972-73 periods are caused by the difference between the negative *ex ante* coefficient and the positive *ex post* coefficient. For example, the 1971 white *ex ante* coefficient value of -.1339 implies that prospective union joiners have a wage level that is $e^{-.1339}-1$ or 12.5 percentage points lower than similar workers in the nonunion base group (i.e., a negative wage gap). The 1972 *ex post* coefficient of .1616 (Table 4) suggests that the union joiner's wage level is $e^{.1616}-1$ or 17.5 percentage points larger than the base group's, after entry into the union sector (i.e, a positive union differential). The black 1971 *ex ante* coefficient of -.2282 and the 1972 *ex post* coefficient of .1962 have similar interpretations. Thus, female union joiners experience a positive wage change as they proceed from a negative *ex ante* wage gap to positive *ex post* union differential. In Figure 3, the 1972-73 u_{01m} wage level estimates for whites and blacks have values that are very similar to the 1971-72 period. The 1970-71 u_{01m} wage level estimates have a similar pattern, but only the white *ex post* union differential is significant. The positive *ex post* union differentials are anticipated since union workers generally receive a positive wag gap relative to similar nonunion cohorts. The negative *ex ante* differentials are not surprising since workers with low wages have the greatest incentive to change jobs.

In the 1973-77 period, the union joiner wage change continues to be positive and at a magnitude similar to the 1970-71 through 1972-73 periods. As shown in Figure 3, the underlying pattern of the u_{01m} wage level coefficients begin to change. By 1973-77, white females no longer have a significant, positive *ex post* union differential (Table 4) and experience a positive union effect only by eliminating the negative *ex ante* wage gap. By the 1977-78 period, neither the white nor black union joiners receive a positive *ex post* union differential, and both female groups receive a positive union joiner effect only from the elimination of a negative 1977 *ex ante* wage gap. In 1978-80, the wage change for white union joiners disappears as both the *ex ante* and *ex post* coefficients are insignificant; and by 1980-82, both white and black union joiner effects are also insignificant since neither group has a sufficiently large *ex ante* or *ex post* wage gap to cause a wage change. The results for the 1973-77 to 1980-82 longitudinal periods suggest a significant weakening of unions' ability to impact female wages at entry level positions. This result is particularly surprising in the recessionary 1978-80 and 1980-82 periods because the union wage gap historically increases during periods of economic downturn.

With respect to male u_{01m} workers, the wage change parameter estimates (Table 3) suggest that both white and black males experience a significant and positive union joiner effect in almost all of the time periods; only the 1980-81 estimate for blacks is insignificant. The white union joiner effect is larger than that for blacks in the 1970-1971, 1971-1976, and 1980-81 periods, which correspond to relatively high or increasing employment. The black union joiner effect is larger in the 1969-70 and 1976-78 periods, which correspond to low or declining unemployment. The 1978-80 period is inconsistent since a large black union effect occurs during a recession, but the unemployment rate was declining by the fourth quarter of 1980. The time trend in Figure 2 suggests that white union joiner effect is procyclical while the black union joiner effect is countercyclical. The female union effect in the 1970-71 through 1973-77 time periods is larger than the white male union effect; and the white male effect is larger, on average, than the black male union effect. The data do not allow for a good comparison since the NLS young men's survey does not cover the 1971-73 time period. The main distinction between the female and male union effect is that the female union effect declines significantly after the 1977-78 period while the magnitude of the male union effect is slightly larger, on average, than the previous periods.

The underlying wage level estimates for male u_{01m} workers (Table 5) suggest that the positive white union effect in the 1969-70, 1970-71, and 1971-76 periods is caused by going from a negative wage gap (relative to the nonunion stayer base group) for prospective union joiners to a positive *ex post* union differential. As illustrated in Figure 4, the trend in the white male union joiner wage level estimates is similar to the trend for females, but the white *ex post* union differential is consistently smaller. Black union joiners, in this period, have insignificant *ex ante* coefficients which suggests that there is no difference between theirs and the base group's wage level when the prospective union joiners are in the nonunion sector. In the *ex post* periods, blacks receive a positive wage change via their large union differential. Because white union joiners go from a negative wage gap in the *ex ante* period to a positive union differential in the *ex post* period, they generally experience a larger percentage wage change than black union joiners (even though blacks receive consistently larger *ex post* union differentials).

As do the females, white and black males experience a change in the pattern of their *ex ante* and *ex post* wage level estimates in the late 1970s and early 1980s. The critical question is why do males continue to have a large positive wage change upon entry into the union sector in this period while females have a reduced or no positive wage effect? Table 5 shows that after 1976 all white *ex post* u_{01m} coefficients are statistically insignificant. White males continue to receive a large positive union joiner effect due to a numerically large negative *ex ante* nonunion differential. The magnitude of this term increases during recessions. In Figure 4, the time tend shows that this coefficient is responsible for the procyclical pattern of the white union joiner effect. Conversely, the black male *ex post* union differential increases, especially during the recessionary years of 1980 and 1981, which gives rise to the countercyclical pattern for the black male union joiner effect. The black *ex ante* u_{01m} coefficient (as opposed to the white nonunion differential) is statistically insignificant and remains at near zero values over the time trend. The exception is the 1980 estimate which has a significant, positive value and causes a decline in the 1980-81 black union effect. Moreover, the time trend (Figure 4) in the *ex ante* wages of new union hires shows that over all time periods white males have lower nonunion wage profile than blacks, which is also the primary reason why whites receive a larger wage change from entry into the union sector.

A similar case can be made for white and black females in the late 1970s and early 1980s since the loss of the negative *ex ante* wage gap is one of the reasons the female union joiner effect disappears in this period. The conflicting *ex ante* wage patterns associated with union entry has interesting implications. Union employers hire low wage nonunion white males (relative to their base group) while they simultaneously hire relatively high wage nonunion females and black males. This pattern is even more dominate during recessions since the white male *ex ante* wage gap moves procyclically. Either workers in the white nonunion base group have greater unobservable productivity characteristics (which cannot be controlled for in the cross-sectional wage equation) and therefore cause the negative *ex ante* wage gap, or union employers are discriminating among equally productive workers by hiring relative more productive, high wage nonunion females and black males while they are willing to hire relatively less productive, low wage nonunion whites.

Union leavers, in theory, should suffer a wage loss when leaving the union sector; that is, experience a negative union effect. The female wage change results (u_{10m} coefficients, Table 2) suggest that the wage loss is procyclical. The largest female wage losses (with the exception of blacks in 1978-80) occur in the 1973-77, 1978-80, and 1980-82 periods which correspond to transitions into or out of economic downturns. In the other time periods, there are no significant wage change estimates (with the exception of whites in 1970-71) which suggests that leaving the union sector in periods of low unemployment fails to have a negative impact on wages. The wage level results (u_{10m} coefficient, Table 4) suggest that the negative wage loss from exiting the union sector during a recession is caused by a procyclical decline in the *ex post* nonunion wage rate. The reason union leavers do not incur a negative wage change during economic upturns is that the *ex post* nonunion wage is relatively higher in these periods. This trend is illustrated in Figure 5. The 1971 white *ex post* coefficient is inconsistent with the overall trend, but it may reflect lack of experience since these workers are very young. The *ex ante* union differential apparently has no significant impact on the wage loss since no *ex ante* u_{10m} coefficients in Table 4 are significant. This also suggests that part of the reason these workers leave the union sector is low relative wages.

With respect to the male cohorts, the wage change results (u_{10m} coefficients in Table 3) suggest that the wage loss for both blacks and whites from leaving the union sector is also procyclical. Blacks appear to suffer the greatest wage loss, at least in the later years (Figure 2). The 1969-70 white male wage change is inconsistent with this argument, but it occurs when most of the cohorts are young, and the lack of experience may cause them to have low nonunion wages in a period of low unemployment. The wage level estimates for union leavers (u_{10m} coefficients in Table 5) suggest that the black *ex ante* union differential has a countercyclical trend which apparently causes the larger union leaver wage loss during recessions. Figure 6 shows that blacks receive relatively large *ex post* union differentials in the high unemployment years of 1976 and 1980; and in 1978, an expansionary year, the union wage gap is small and insignificant. The *ex post* nonunion differential appears to have a procyclical time trend; but none of the parameter estimates in Table 5 is significant. In the early 1970s, white males have large union leaver wage losses resulting from going from a positive *ex ante* union differential to a large negative *ex post* nonunion differential. It is likely that the low nonunion wage structure is due to the inexperience of white union leavers in this period. In the later 1970s when whites have a smaller procyclical union leaver wage loss, the wage change is caused by a large negative *ex post* nonunion wage gap that appears to increase (in absolute value) during recessions. The white *ex ante* union differential has little impact on wages (the pattern is similar for white females), and the 1978 estimate is negative and significant. Again, the inability of unions to capture a positive union differential likely causes these workers to exit the union sector.

The wage level parameter estimates for union stayers capture the wage effect of union seniority ladders. The estimates for u_{11t} workers (females, Table 4; and males, Table 5) imply that both females and males receive a significant union wage premium in nearly all periods (only the white coefficient in the 1980 *ex post* period is insignificant).[10] The u_{11t} estimates imply that females and black males generally receive a larger union premium than white males. In the early 1970s, the females' union premium is essentially equal to the white males', but this is probably due to the relative lack of female seniority. By the mid-1970s, females have a significantly larger union differential. The most striking trend (Figures 7 and 8) is that the union differential for white males and white females declines from the mid-1970s. By the early 1980s, the female union stayer differential is below 15 percentage points and the white differential is between 5 percentage points and zero. The black female and black male union differential remains around 20 percentage points and retains the conventional countercyclical pattern. The white differentials no longer move in a countercyclical pattern. This attenuation of the white union stayer wage gap is consistent with the decline in the union differentials associated with the union joiner and leaver groups.

Tables 6 and 7 represent a summary of the "true" return for female and male union joiners, respectively. The female true union effect (columns 3 and 4) has a similar pattern to the unadjusted union joiner effect (Figures 1 and 2), but is generally smaller due to the mobility adjustment. In the 1970-71 through 1973-77 periods, the female true union effect (both whites and blacks) is slightly larger, on average, than the white male true union effect for the 1969-70 through 1971-76 periods. The sizable female true union effect is cause by a union differential for new hires (column 1) that is larger, on average, than the males'; and by a large (in absolute value) negative female selectivity-in-hiring component (column 2). In the 1977-78 through 1980-82 periods, the female true union effect shrinks and is much smaller, on average, than the white and black males'. This attenuation is caused by a decline in the female new hire differential and a selectivity term that is now positive (with the exception of the white 1977-78 term). The selectivity term which is based on wages in the nonunion sector can be interpreted as an estimate of relative productivity (at the current job). While the female selectivity term increases in magnitude over time, the white male selectivity component acquires a large negative value in the later periods (-.1653 by the 1980-81 period). Conversely, black union joiners exhibit the expected positive selectivity term (with the exception of 1978-80). Negative selectivity implies that white male union joiners are relatively less productive than the nonunion stayers who search for a new job. Unionized females and black males, however, generally have stronger productivity characteristics than their nonunion counterparts. The negative selectivity result is surprising since most cross-sectional studies suggest that union workers, white and black, are more productive than comparable nonunion workers.[11]

CONCLUSIONS

In general, the results show that unions' ability to impact wages for all white workers and black female union joiners has attenuated over the 1970s to the early 1980s. Union effects for all other black workers over this time period have remained relatively constant. From the early to mid-1970s, the female wage change from entering the union sector is typically larger than the wage change for males. Estimates for the "true" union joiner effect suggest that the female wage change is greater because of a large negative selectivity-in-hiring term (i.e., they are low wage nonunion workers). Therefore, they experience a larger wage change from entering the union sector. In the later 1970s and early 1980s, the female wage change becomes significantly smaller (and in some instances it is actually negative), as the female selectivity term becomes a large positive value (i.e., they are now high wage nonunion workers). While the female union effect declines over time, the magnitude of the male union effect remains fairly stable over time. Union joiner effects are larger for white males primarily because white union joiners have lower productivity characteristics than black male union joiners, and therefore receive a larger wage change from entering the union sector. The larger white union effect generally does not result from a higher relative wage distribution in the union sector. The results suggest that either union employers are willing to hire less productive white males from the labor pool, or only low wage white males are willing to enter the union sector. The results also suggest that union employers may discriminate against females and blacks by hiring only the best qualified of these workers.

The union stayer estimates (u_{11s}) reflect the relative wage gains from seniority ladders. The results show that black females and black males receive the greatest relative benefits from remaining in the union sector. Both the white females' and the white males' union stayer wage differential declines over the mid-1970s to early 1980s. The decline in the white male wage gap is so great that difference between union and nonunion wage rate is less than 5 percentage points in the 1980s.

NOTES

1. To make a more direct comparison between similar union and nonunion workers, all agriculture and self-employed cohorts are eliminated from the sample.

2. The exact length of time over each longitudinal period vary due to variations in the number of years between surveys which report union status. The longitudinal periods vary from one to five years. It is unlikely

that the difference longitudinal lengths will bias the results since the cohorts are of the same age group, and the variation in wages are not subject to demographic tends.

3. Ashenfelter [2] demonstrates that the union differential declines in periods of low unemployment and/or accelerating inflation (and vice versa) as nonunion wages react to prevailing market conditions. The longitudinal periods in this study that coincide with a recession 1973-77, 1978-80, and 1980-82 periods in the female survey; and the 1971-76, 1976-78, and 1980-81 periods in the male survey.

4. The marital categorical variable equals one if married, and the health categorical variable equals one if the interviewee reports that his or her health hinders job performance.

5. This technique for selectivity bias correction was developed by Heckman [15]. The λ_{it} term is estimated from a probit model that predicts inclusion in the sample (i.e., the wage equation). The probit model includes all the explanatory variables in the wage equation with the exception of the union variables, senority, and senority squared which are determined by one's employment decision. Also included are variables that help determine female labor force participation: school enrollment status, number of children, and other household income. The correction for selectivity bias is not used in the male wage equation since most males in the sample are in the labor force, and therefore sample selectivity is not a serious issue.

6. Female sample mean values are reported for only the 1970-1971 time period, and male sample mean values are reported for only the 1969-1971 time period to avoid a cumbersome array of tables. The longitudinal sample reflects the number of observations after attrition due to missing variables in either cross-sectional survey. The 1971-72 female sample has 1120 white observations and 376 black observations; the 1972-73 sample has 1179 white observations and 424 black observations; the 1973-77 sample has 1210 white observations and 493 black observations; the 1977-78 sample contains 1289 white observations and 542 black observations; the 1978-80 sample has 1085 white observations and 459 black observation; and the 1981-82 sample has 1239 white observations and 557 black observations. The 1971-76 male sample has 1526 white observations and 425 black observations; the 1976-78 sample contains 1731 white observations and 478 black observations; the 1978-80 sample has 1738 white observations and 481 black observation; and the 1980-81 sample has 1760 white observations and 522 black observations.

7. The change in a individual's labor force experience over the longitudinal period will be identical for all interviewees. Additionally, the change in education is included in the wage change model. Typically, this variable is assumed to be time invariant. However, there are differences in the highest grade attained across all sampling periods and the variable proves to be significant for females and black males.

8. In practice, the first-difference and wage change estimates will differ due to different error structures.

9. The parameter estimates for the probit model and the explanatory variables in the *ex ante* and *ex post* wage level and wage change equations are available in an unpublished appendix.

10. Since the u_{11s} parameter estimates are from the cross-sectional wage model, they will contain a positive heterogeneity bias.

11. This result is inconsistent with the well established findings of a positive selectivity bias term estimated using a probit model (in particular, see Lee [18]; and Duncan and Leigh [9]).

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Appendix

Tables 1-7

Figures 1-8

Table 1
Union Variable Definitions and Mean Values (Standard Deviation)

Union Variables	Definition	White Females	White Males	Black Females	Black Males
u_{01m}	1 if union joiner who moves	.031 (.030)	.038 (.191)	.041 (.039)	.057 (.232)
u_{10m}	1 if union leaver who moves	.031 (.030)	.045 (.207)	.050 (.047)	.075 (.265)
u_{11m}	1 if union stayer who moves	.032 (.031)	.086 (.282)	.058 (.055)	.129 (.335)
u_{00m}	1 if nonunion stayer who moves	.378 (.235)	.332 (.471)	.368 (.233)	.293 (.456)
u_{01s}	1 if union joiner who stays	.030 (.030)	.021 (.143)	.041 (.039)	.026 (.161)
u_{10s}	1 if union leaver who stays	.023 (.023)	.009 (.096)	.032 (.031)	.034 (.181)
u_{11s}	1 if union stayer who stays	.102 (.092)	.168 (.374)	.122 (.108)	.161 (.368)
u_{00s}	1 if nonunion stayer who stays (base)	.373 (.293)	.301 (.296)	.467 (.499)	.225 (.274)
Sample Size:†		1117	1737	341	529

†The female and male sample size is based on the 1970-71 period.

Table 2
Parameter Estimates (and Standard Errors)
for the Female Wage Change Regressions

Union Variables	White Sample	Black Sample
1970-71 Period:		
u_{01m}	.2198 (.0584)†	.2259 (.0809)†
u_{10m}	-.1074 (.0572)§	.0625 (.0748)
u_{00m}	.0211 (.0301)	.1258 (.0443)†
u_{11s}	.0210 (.0329)	.0248 (.0483)
1971-72 Period:		
u_{01m}	.2439 (.0600)†	.3374 (.0811)†
u_{10m}	-.0431 (.0639)	.0020 (.0862)
u_{00m}	.0728 (.0263)†	.0249 (.0488)*
u_{11s}	.0104 (.0286)	.0396 (.0400)
1972-73 Period:		
u_{01m}	.3733 (.0590)†	.3609 (.0564)†
u_{10m}	.0671 (.0622)	-.0132 (.0817)
u_{00m}	.0269 (.0261)	.0989 (.0384)†
u_{11s}	-.0083 (.0266)	-.0073 (.0342)
1973-77 Period:		
u_{01m}	.1656 (.0514)†	.2611 (.0725)†
u_{10m}	-.2072 (.0535)†	-.1652 (.0705)†
u_{00m}	.0125 (.0312)	.0356 (.0456)
u_{11s}	-.0037 (.0433)	.0735 (.0589)
1977-78 Period:		
u_{01m}	.2246 (.0458)†	.0838 (.0680)*
u_{10m}	-.0447 (.0523)	-.0042 (.0849)
u_{00m}	.0232 (.0231)	.0970 (.0393)†
u_{11s}	-.0189 (.0239)	-.0035 (.0384)
1978-80 Period:		
u_{01m}	.0139 (.0455)	.1680 (.0606)†
u_{10m}	-.2153 (.0598)†	.0404 (.1134)
u_{00m}	.0284 (.0263)*	.0838 (.0493)§
u_{11s}	-.0104 (.0288)	.0111 (.0399)
1980-82 Period:		
u_{01m}	.0260 (.0624)	-.0342 (.0694)
u_{10m}	-.1477 (.0450)†	-.1414 (.0553)†
u_{00m}	-.0385 (.0237)*	-.0045 (.0354)
u_{11s}	.0346 (.0292)	-.0028 (.0341)

*A .20 probability value of a nonzero estimate--two tail test.
† $p \leq .01$
‡ $p \leq .05$
§ $p \leq .10$

Table 3
Parameter Estimates (and Standard Errors)
for the Male Wage Change Regressions

Union Variables	White Sample	Black Sample
1969-70 Period:		
u_{01m}	.1395 (.0336)†	.1524 (.0443)†
u_{10m}	-.3059 (.0344)†	-.0932 (.0549)§
u_{00m}	.0047 (.0213)	-.0413 (.0360)*
u_{11s}	-.0006 (.0204)	.0386 (.0348)
1970-71 Period:		
u_{01m}	.2196 (.0383)†	.1049 (.0558)§
u_{10m}	-.1499 (.0361)†	-.0487 (.0504)
u_{00m}	.0346 (.0186)§	.0479 (.0339)*
u_{11s}	.0173 (.0208)*	.0539 (.0380)
1971-76 Period:		
u_{01m}	.2435 (.0411)†	.1166 (.0678)§
u_{10m}	-.1718 (.0471)†	-.1995 (.0734)†
u_{00m}	.0498 (.2099)§	.0732 (.0522)*
u_{11s}	-.0417 (.0310)*	-.0042 (.0534)
1976-78 Period:		
u_{01m}	.2111 (.0423)†	.2527 (.0628)†
u_{10m}	-.0188 (.0484)	-.0675 (.0960)
u_{00m}	.0654 (.0239)†	.1727 (.0529)†
u_{11s}	-.0260 (.0190)*	-.0647 (.0372)§
1978-80 Period:		
u_{01m}	.1542 (.0491)†	.2515 (.1002)†
u_{10m}	-.0415 (.0479)	-.1532 (.0886)§
u_{00m}	.0185 (.0247)	.0149 (.0510)
u_{11s}	-.0357 (.0191)§	-.0112 (.0405)
1980-81 Period:		
u_{01m}	.2911 (.0671)†	.1264 (.1172)
u_{10m}	-.1953 (.0506)†	-.3281 (.0985)†
u_{00m}	-.0134 (.0205)	-.0432 (.0499)
u_{11s}	-.0106 (.0170)	.0396 (.0374)

*A .20 probability value of a nonzero estimate–two tail test.
† $p \leq .01$
‡ $p \leq .05$
§ $p \leq .10$

Table 4
Parameter Estimates (and Standard Errors)
for the Female Cross-Sectional Wage Level Regressions

Union Variables	*(ex ante)* Period		*(ex post)* Period	
	White Sample	Black Sample	White Sample	Black Sample
1970-71 Period:				
u_{01m}	-.0419 (.0575)	-.0884 (.0855)	.1842 (.0661)†	.0990 (.0973)
u_{10m}	-.0049 (.0570)	-.0006 (.0801)	-.1131 (.0666)‡	.0968 (.0927)
u_{00m}	-.1295 (.0244)†	-.1200 (.0435)†	-.0849 (.0405)‡	-.0088 (.0635)
u_{11s}	.1380 (.0340)†	.0947 (.0458)§	.1575 (.0353)†	.1493 (.0590)†
1971-72 Period:				
u_{01m}	-.1339 (.0639)‡	-.2282 (.0862)†	.1616 (.0662)†	.1962 (.0899)‡
u_{10m}	-.0211 (.0682)	-.1066 (.0929)	-.0104 (.0710)	.0223 (.0999)
u_{00m}	-.1012 (.0258)†	-.0382 (.0440)	.0068 (.0340)	.0699 (.0526)*
u_{11s}	.1920 (.0316)†	.1250 (.0445)†	.2034 (.0310)†	.1689 (.0444)†
1972-73 Period:				
u_{01m}	-.2229 (.0626)†	-.1707 (.0633)†	.1722 (.0682)†	.2870 (.0695)†
u_{10m}	-.0584 (.0657)	.0824 (.0877)	.0448 (.0711)	.1463 (.0919)*
u_{00m}	-.0577 (.0234)†	-.1279 (.0389)†	.0083 (.0326)	.0565 (.0501)
u_{11s}	.2173 (.0299)†	.1659 (.0411)†	.2018 (.0309)†	.1602 (.0402)†
1973-77 Period:				
u_{01m}	-.1634 (.0465)†	-.1509 (.0578)†	.0209 (.0514)	.1283 (.0724)§
u_{10m}	.0702 (.0456)*	.0535 (.0552)	-.1248 (.0520)†	-.0592 (.0701)
u_{00m}	-.1145 (.0253)†	-.1024 (.0354)†	-.0834 (.0328)†	-.0363 (.0485)
u_{11s}	.1387 (.0411)†	.1570 (.0497)†	.1384 (.0430)†	.2273 (.0568)†
1977-78 Period:				
u_{01m}	-.1735 (.0542)†	-.1437 (.0672)‡	.0397 (.0538)	-.1109 (.0691)§
u_{10m}	-.0280 (.0591)	.0827 (.0848)	-.0354 (.0602)	.0176 (.0862)
u_{00m}	-.0897 (.0251)†	-.1573 (.0381)†	-.0385 (.0299)*	-.1042 (.0418)†
u_{11s}	.1415 (.0292)†	.1916 (.0413)†	.1315 (.0283)†	.1833 (.0403)†
1978-80 Period:				
u_{01m}	-.0524 (.0487)	.0345 (.0630)	-.0371 (.0579)	.1639 (.0686)†
u_{10m}	.0150 (.0627)	-.1043 (.1168)	-.1585 (.0713)‡	-.1026 (.1169)
u_{00m}	-.1118 (.0275)†	-.0464 (.0469)	-.0771 (.0381)‡	.0079 (.0594)
u_{11s}	.1054 (.0326)†	.2319 (.0432)†	.0985 (.0353)†	.2099 (.0413)†
1980-82 Period:				
u_{01m}	.0443 (.0687)	-.0192 (.0796)	.1041 (.0754)*	-.0525 (.0672)
u_{10m}	-.0195 (.0502)	.0783 (.0634)	-.1622 (.0554)†	-.0356 (.0672)
u_{00m}	-.1078 (.0260)†	-.0297 (.0391)	-.1286 (.0318)†	-.0146 (.0445)
u_{11s}	.0672 (.0330)‡	.1515 (.0409)†	.0900 (.0353)†	.1325 (.0414)†

*A .20 probability value of a nonzero estimate--two tail test.
$^\dagger p \leq .01$
$^\ddagger p \leq .05$
$^\S p \leq .10$

Table 5
Parameter Estimates (and Standard Errors)
for the Male Cross-Sectional Wage Level Regressions

Union Variables	*(ex ante)* Period White Sample	*(ex ante)* Period Black Sample	*(ex post)* Period White Sample	*(ex post)* Period Black Sample
1969-70 Period:				
u_{01m}	-.0886 (.0347)†	-.0266 (.0437)	.0604 (.0409)*	.1428 (.0551)†
u_{10m}	.1153 (.0352)†	.1612 (.0577)†	-.1732 (.0416)†	.0681 (.0620)
u_{00m}	-.1428 (.0209)†	-.0361 (.0354)	-.1210 (.0277)†	-.0745 (.0423)§
u_{11s}	.1246 (.0218)†	.1888 (.0393)†	.1258 (.0235)†	.2182 (.0399)†
1970-71 Period:				
u_{01m}	-.1607 (.0464)†	-.0411 (.0625)	.0436 (.0500)	.0844 (.0706)
u_{10m}	.0145 (.0427)	.0937 (.0557)§	-.1483 (.0468)†	.0395 (.0629)
u_{00m}	-.1118 (.0243)†	-.0731 (.0397)§	-.0848 (.0287)†	-.0186 (.0451)
u_{11s}	.1111 (.0253)†	.1811 (.0443)†	.1347 (.0251)†	.2133 (.0462)†
1971-76 Period:				
u_{01m}	-.1308 (.0370)†	-.0168 (.0572)	.1059 (.0408)†	.1468 (.0725)‡
u_{10m}	.0852 (.0410)‡	.0888 (.0596)*	-.0548 (.0454)	-.0365 (.0766)
u_{00m}	-.0744 (.0243)†	-.0875 (.0419)‡	.0007 (.0313)	-.0023 (.0571)
u_{11s}	.1264 (.0290)†	.1553 (.0484)†	.0738 (.0290)†	.1543 (.0558)†
1976-78 Period:				
u_{01m}	-.2156 (.0455)†	-.0274 (.0613)	-.0338 (.0580)	.1556 (.0778)‡
u_{10m}	-.0359 (.0519)	.1903 (.0962)‡	-.0799 (.0653)	.1161 (.1087)
u_{00m}	-.1167 (.0241)†	-.0390 (.0525)	-.0805 (.0386)‡	.0635 (.0691)
u_{11s}	.0684 (.0218)†	.2295 (.0402)†	.0398 (.0249)*	.1545 (.0430)†
1978-80 Period:				
u_{01m}	-.1227 (.0598)‡	-.0354 (.1010)	.0388 (.0629)	.2870 (.1165)†
u_{10m}	-.1230 (.0555)‡	.0577 (.0902)	-.1662 (.0585)†	-.0690 (.1035)
u_{00m}	-.0666 (.0281)‡	-.0208 (.0506)	-.0394 (.0346)	.0677 (.0655)
u_{11s}	.0639 (.0242)†	.1703 (.0427)†	.0277 (.0238)	.1581 (.0454)†
1980-81 Period:				
u_{01m}	-.2428 (.0897)†	.1770 (.1309)*	.0561 (.0940)	.2851 (.1266)‡
u_{10m}	.0536 (.0648)	.2234 (.1012)‡	-.1333 (.0704)§	-.0409 (.1008)
u_{00m}	-.0775 (.0313)†	-.0687 (.0553)	-.1067 (.0372)†	-.1112 (.0578)‡
u_{11s}	.0348 (.0236)*	.2264 (.0435)†	.0477 (.0243)‡	.2699 (.0413)†

*A .20 probability value of a nonzero estimate--two tail test.
† $p \leq .01$
‡ $p \leq .05$
§ $p \leq .10$

Table 6
Summary of the Net Union Joiner Wage Effect for the Female Wage Level and Wage Change Models

	Union Wage Differentials New Hires $(\delta_{01} - \delta_{00})_2$	Selectivity in Hiring $(\delta_{01} - \delta_{00})_1$	True Union Wage Differential Net of Selectivity (1)-(2)	True Wage Change of Union Joiners Net of Mobility $(d_{01} - d_{00})$	Union Stayer Union Wage Differential *ex post*
	(1)	(2)	(3)	(4)	(5)
1970-71 Period:					
White:	.2691	.0876	.1815	.1987	.1575
Black:	.1078	.0316	.0762	.1001	.1493
1971-72 Period:					
White:	.1548	-.0327	.1875	.1711	.2034
Black:	.1263	-.1900	.3163	.3125	.1689
1972-73 Period:					
White:	.1639	-.1652	.3291	.3464	.2018
Black:	.2305	-.0428	.2733	.2620	.1603
1973-77 Period:					
White:	.1043	-.0489	.1532	.1531	.1383
Black:	.1646	-.0485	.2131	.2255	.2273
1977-78 Period:					
White:	.0782	-.0838	.1620	.2014	.1315
Black:	.0067	.0136	-.0069	-.0132	.1833
1978-80 Period:					
White:	.0400	.0594	-.0194	-.0145	.0984
Black:	.1560	.0809	.0751	.0842	.2099
1980-82 Period:					
White:	.2327	.1521	.0806	.0645	.0900
Black:	-.0379	.0105	-.0484	-.0297	.1325

Table 7
Summary of the Net Union Joiner Wage Effect for the Wage Level and Wage Change Models

	Union Wage Differentials New Hires $(\delta_{01} - \delta_{00})_2$	Selectivity in Hiring $(\delta_{01} - \delta_{00})_1$	True Union Wage Differential Net of Selectivity (1)-(2)	True Wage Change of Union Joiners Net of Mobility $(d_{01} - d_{00})$	Union Stayer Union Wage Differential *ex post*
	(1)	(2)	(3)	(4)	(5)
1969-70 Period:					
White:	.1814	.0542	.1272	.1348	.1258
Black:	.1963	.0050	.1913	.1937	.2182
1970-71 Period:					
White:	.1284	-.0489	.1773	.1850	.1640
Black:	.1030	.0321	.0709	.0570	.2389
1971-76 Period:					
White:	.1052	-.0564	.1616	.1937	.0738
Black:	.1491	.0707	.0784	.0434	.1543
1976-78 Period:					
White:	.0467	-.0991	.1458	.1457	.0398
Black:	.0921	.0116	.0805	.0800	.1545
1978-80 Period:					
White:	.0782	-.0561	.1343	.1357	.0277
Black:	.2193	-.0149	.2342	.2365	.1581
1980-81 Period:					
White:	.1629	-.1653	.3281	.3045	.0477
Black:	.3963	.2457	.1506	.1696	.2699

FIGURE 1

Female Union Joiner and Leaver Wage Change Estimates

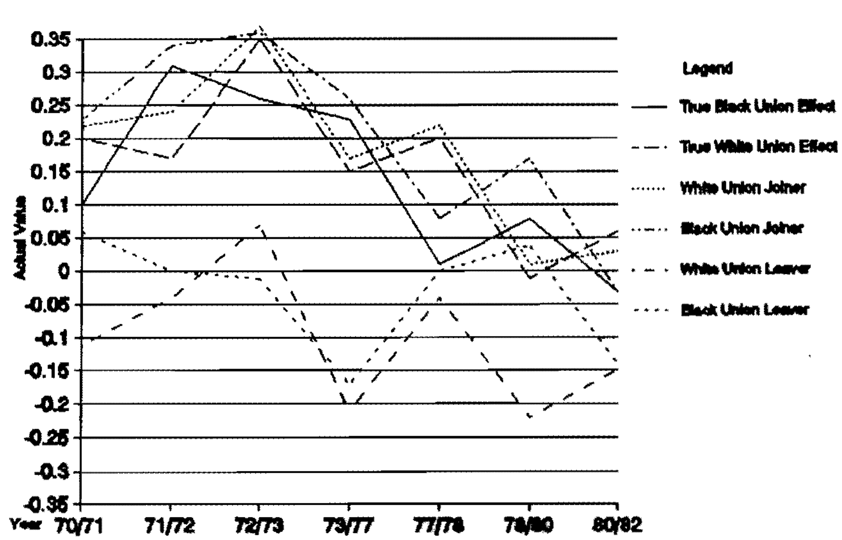

FIGURE 2

Male Union Joiner and Leaver Wage Change Estimates

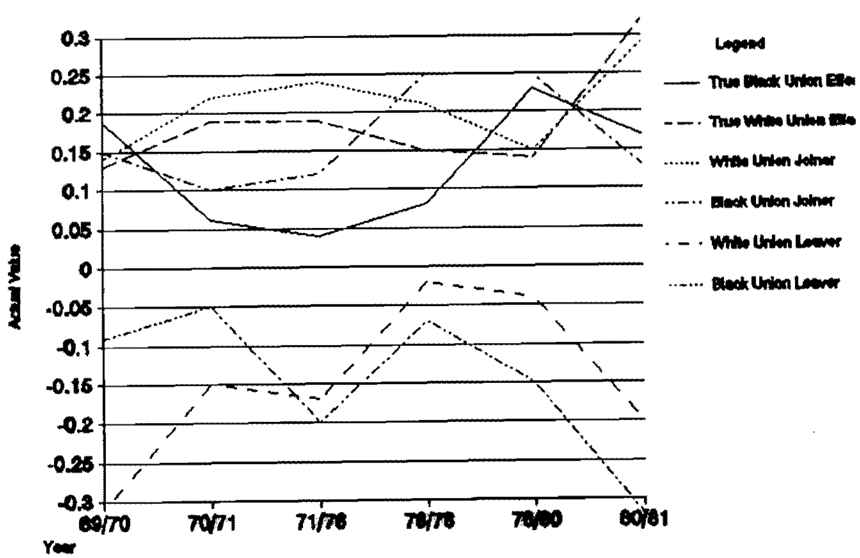

FIGURE 3

Female Union Joiner Wage Level Estimates

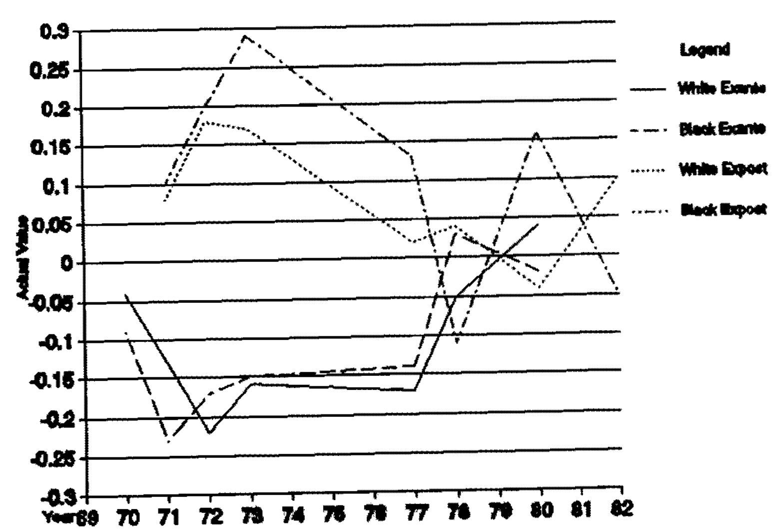

FIGURE 4

Female Union Stayer Wage Level Estimates

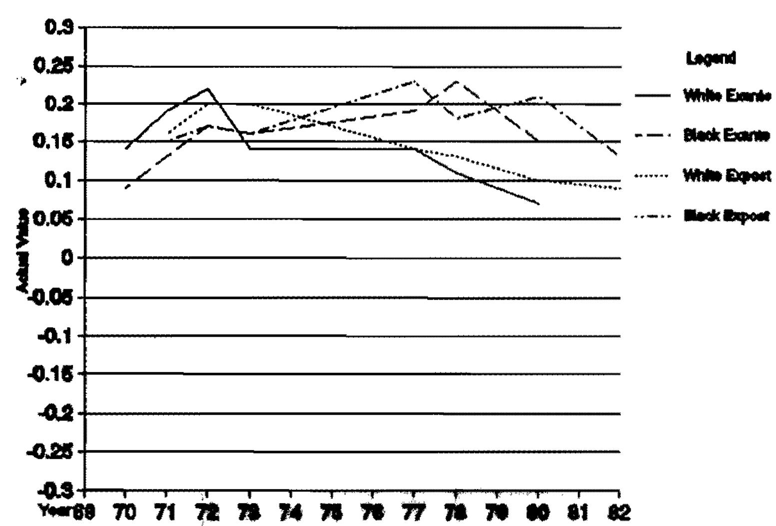

FIGURE 5

Female Union Leaver Wage Level Estimates

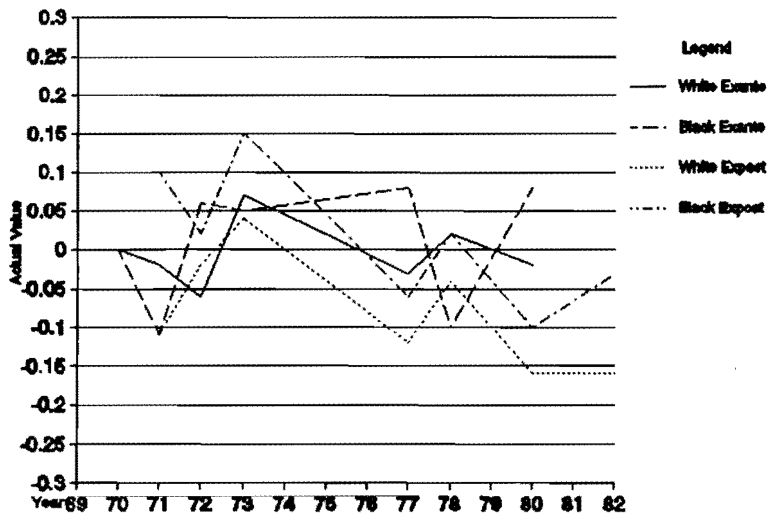

FIGURE 7

Male Union Leaver Wage Level Estimates

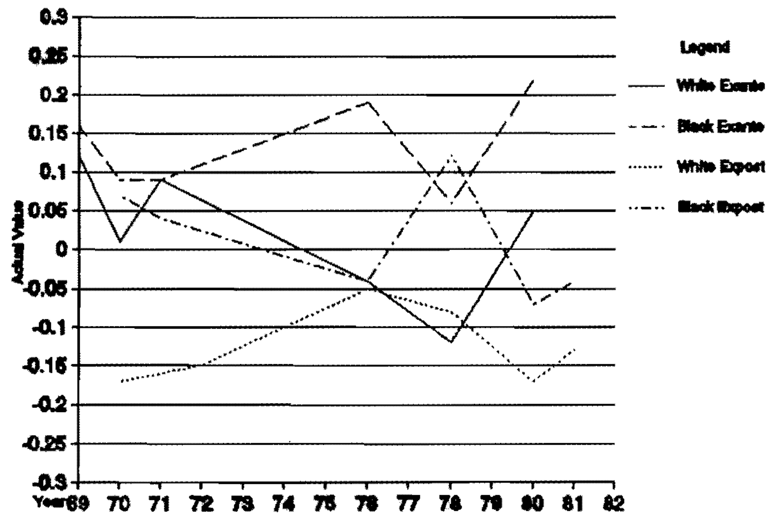

FIGURE 6

Male Union Joiner Wage Level Estimates

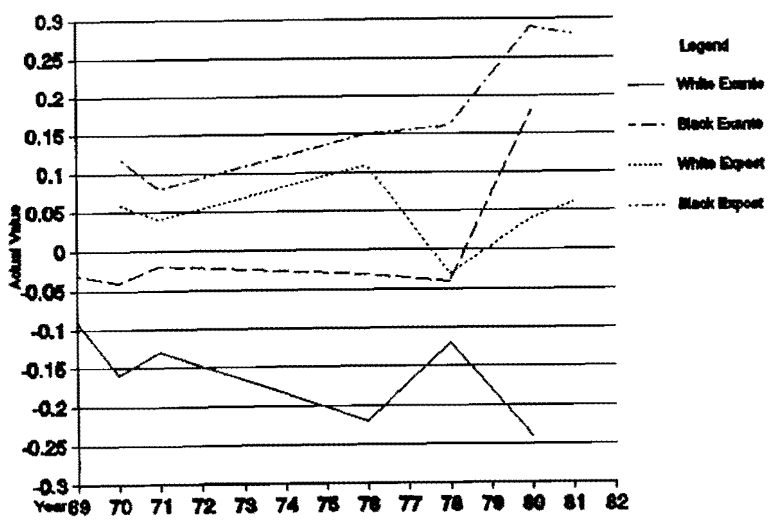

FIGURE 8

Male Union Stayer Estimates