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MONOPSONISTIC DISCRIMINATION AND SEX DIFFERENCES IN WAGES

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Introduction

Wage differences between men and women have often been attributed to three general causes: first, the differences in personal characteristics such as education or experience; second, the psychic cost of working with women, which we call "taste discrimination" and is discussed, for example, by Becker (1957) and Arrow (1972); and third, women's relative immobility resulting in a poor bargaining position and monopsonistic exploitation in the labor market as suggested by Krueger (1963) and Thurow (1968). In fact, since monopsony could affect both men and women, this third cause allows for the possibility that men are more hurt by monopsonistic discrimination than are women. It is this last type of wage discrimination that has received the least attention in the literature and that we plan to focus on here.

Many people have discussed the interdependencies of these sources of wage differentials. The first two are certainly related. For instance, it is frequently argued that the amount of schooling that women choose to obtain or the experience they have is affected by the discrimination they face in the labor market. In addition, taste and monopsonistic discrimination are mutually dependent. Both social mores and alternative employment opportunities are affected by taste discrimination and as well affect the likelihood of monopsony. For instance, the relative immobility of women, which is a prerequisite for monopsony, depends on both social customs and on employment opportunities in other sectors of the economy. Two causes of monopsonistic discrimination can thus be distinguished.

Several authors have noted the importance of socialization in explaining female labor force behavior. On the one hand, women might have a preference for working close to home or in the same area as their husbands. Malkiel and Malkiel (1973) for instance, write "it is generally presumed by employers that a working wife will quit her job when the husband's job changes so as to require the family to move."[1] By extension, once the husband has found a job, the wife typically looks for employment within a geographically restricted area. Her geographical immobility then might result in a lower rate of return to her skills because of the limited selection of jobs. But as well, she might be subject to monopsonistic discrimination due to her immobility. Alternatively, women might prefer certain occupations over others. Oaxaca (1973) argues that, "social conditioning starting with childhood experiences is in large part responsible for the seemingly voluntary occupational choices of so many women."[2] Again the choice of jobs is limited, resulting in a lower return to women's skills as well as decreasing a woman's occupational mobility.

Quite a different explanation of monopsony comes not from women's personal preferences but from the industrial structure. Women might be excluded from some occupations and industries, as a result of discrimination by employers and employees. As Bergmann (1971) argues, they are then crowded into the remaining sectors of the economy. Moreover, they become effectively less mobile within the occupational structure. Once again, we could expect to find a lower rate of return to women's skills for two reasons: first, the limited choice of jobs and second, the additional effect of monopsony.

This discussion should make clear that it is impossible, both theoretically and empirically, to disentangle the separate effects of taste or monopsonistic discrimination on characteristic rates of return. Our aim in this paper, however, is simply to test for the existence of monopsony and to measure the overall impact on wages due to differentials in rates of return and characteristics.

In Section I we review briefly the existing theoretical and empirical work on monopsony in the labor market. The theoretical implications of monopsonistic discrimination are analyzed under alternative supply elasticity assumptions and according to the degree of exploitation in Section II, in which it is found that under certain conditions an affirmative action wage equalization policy may hurt women. Section III presents our empirical methodology for testing for monopsony and a statistical analysis based on the 1970 Public Use Samples. The results obtained indicate that the wages of never-married women are significantly affected by monopsony in the labor market, whereas men's wages do not appear to be. We conclude that monopsonistic discrimination thus accounts for some proportion of the difference between male and female wages that has been formerly attributed solely to taste discrimination.

I. The Literature

Monopsony is a theme that appears repeatedly in the literature in general terms, but rarely with any empirical content. The standard theoretical references are still Robinson (1932), Rothschild (1954), Krueger (1963), and Bronfenbrenner (1956). While Bronfenbrenner considers a non-profit maximizing firm, Robinson, Rothschild and Krueger take a more traditional approach, which we will also follow. The second section of this paper describes in some detail our theoretical framework.

The little empirical work that has been done is indirect and only loosely tied to the theory. Douglas (1939, 1948), for instance, finds that actual factor shares are not significantly different from competitively determined shares, but he does not distinguish between monopolistic and monopsonistic exploitation.

Bunting (1962) measures concentration ratios of employers in 1774 areas, but does not test how concentration affects wages. Finally, Nelson (1973) tries to measure the elasticity of labor supply to particular industries, but again, does not relate the elasticity to actual wage determination. None of the empirical work looks into the possibility of monopsonistic discrimination, although Robinson poses the problem and the standard literature on discrimination pays attention to it.[3]

II. Theoretical Framework

While our theoretical framework is not original, there are several assumptions and implications which we want to emphasize. From this discussion, we obtain several hypotheses which we test in Section III.

We distinguish several cases according to the degree of monopsony power held by an employer[4] and to the supply elasticity of labor. On the one hand, women might be more subject to monopsony power than are men; but on the other hand, they might have greater supply elasticities as a result of their alternatives in household production. For each of these cases, we make predictions first about the wages of women relative to men; second about the response of wages to changes in supply elasticities; third about the effect on employment and wages of a wage equalization affirmative action policy; and fourth about the validity of the usual estimates of pure taste discrimination.[5]

The standard theory of monopsonistic discrimination considers only the case where both groups face a monopsonist and both groups have positive (or infinite) supply elasticities.[6] The basic result is well known: the group with a lower supply elasticity is paid a lower wage. In addition, we can say that an increase in the supply elasticity will increase the wage; a wage equalization policy will lower or leave unchanged (raise) the wage of the higher (lower) paid group and lower or leave unchanged (raise) its employment;[7] and finally, the usual measure of taste discrimination is an underestimate if the group affected by taste discrimination has the higher supply elasticity.

Each of these propositions can be explained with the following model of a discriminating monopsonist:

The monopsonist maximizes profits by hiring either female or male labor, which are perfect substitutes in production but which have different supply functions.

$$\max \pi = R(L_f + L_m) - W_f(L_f)L_f - W_m(L_m)L_m$$

w.r.t. L_f, L_m

giving the first order conditions:

$$MRP\ (L_f + L_m) - W_f\left(1 + \frac{dW_f}{dL_f}\ \frac{L_f}{W_f}\right) = MRP - W_f\ \left(1 + \frac{1}{\eta_f}\right) = 0$$

$$MRP\ (L_w + L_m) - W_m\left(1 + \frac{dW_m}{dL_m}\ \frac{L_m}{W_m}\right) = MRP - W_m\ \left(1 + \frac{1}{\eta_m}\right) = 0$$

where MRP is: marginal revenue product

W_i: the wage of group i, f = women, m = men

L_i: the employment of group i

η_i: the supply elasticity of group i

Our first conclusion follows from the first order conditions. The second comes from differentiating a first order condition, with respect to η, holding MRP constant. The third can be obtained by letting profits be a function of W_f and k, where $kW_f = W_m$. The second order conditions then imply that $dW_f/dk < 0$. If $W_f > W_m$ ($W_f<W_m$) in the case of a discriminating monopsonist, then a wage equalization policy by changing the value of k will result in a decrease (increase) in W_f.[8] The wage changes then imply changes in employment along the relevant supply curves. Finally, if women have higher supply elasticities than men, such that the discriminating monopsonist tends to pay women a higher wage, then the observed wage differential for standardized labor quality is an underestimate of the effects of taste discrimination. On the other hand, if men are perfectly mobile among firms, the discriminating monopsonist pays men the competitive wage which is above the women's wage. Then the observed wage differential is an overestimate of the effects of taste discrimination.

Figures 1 and 2 illustrate these cases.

In figure 1, we take the case of a pure monopsonist which has some control over the wage paid to men as well as to women. In order to maximize profits, the discriminating monopsonist will hire L_{fo} women and L_{mo} men, paying wages equal to W_{fo} and W_{mo}. Since the supply elasticity for women exceeds that of men, by assumption, women's wage exceeds that of men. From this, we can see that an increase in the supply elasticity will result in a higher wage paid. Moreover, if women are typically paid less then men, the effects of taste discrimination must more than offset the effects of monopsonistic discrimination (given our assumption about the relative supply elasticities). Under a wage equalization affirmative action plan, the employer no longer equalizes the MFC across groups. Instead, the employer will hire workers up to the point where the curve which is marginal to the aggregate supply function equals marginal revenue product. In figure 1, the nondiscriminating monopsonist will employ L_{f1} women and L_{m1} men, decreasing wages and employment of women and increasing wages and employment of men, compared with the discriminating monopsonist.

Figure 2 illustrates the case where the employer is a monopsonist in the employment of women, but a perfectly competitive employer of men. Since the monopsonist pays women a lower wage then men, the observed wage differential is partially attributable to monopsony and only partially attributable to taste discrimination. That is, the usual estimates of taste discrimination are overestimates.

Once we consider the case of negative supply elasticities, our results change. A profit maximizing monopsonist will never operate along the negative region of a labor supply curve. An equilibrium where marginal factor costs equals marginal revenue product at a point where the supply elasticity is negative might be a local, but not a global maximum. Therefore, negative labor supply elasticities are consistent only with employers operating in a competitive labor market.

III. Testing for Monopsony

In this section we describe our test for the existence of monopsony in male and female labor markets. Our procedure, which is based on the main implication of the theory outlined in Section II--that if markets are characterized by monopsony, supply elasticities should be significantly and positively related to wage rates--is to estimate the association between wages and supply elasticities, controlling for personal characteristics. The empirical analysis will in addition enable us to ascertain whether a simple "affirmative action" wage equalization policy is likely to benefit or harm women and whether the usual residual measures of taste discrimination against women are under or overestimates of the "true" discrimination component of male and female wage differentials.

Our procedure for testing for the predicted positive relationship between wages paid and the supply elasticities of men and women consists of two stages: in the first, samples of males and females are divided into geographical groups corresponding to identifiable labor markets. Supply elasticities of the individuals within each market for each sex are estimated. In the second stage these elasticity estimates based on the individuals in each area are then entered into an earnings equation which is run across the set of markets. Thus, while the theory of monopsony is formulated in terms of the supply elasticities to individual firms, our empirical analysis is based on relations between labor market aggregates. Therefore, for the test outlined to be valid it is necessary to assume that the strength of the association between aggregate wages and the market elasticities of supply is a continuous positive function of the degree of monopsony characterizing the labor market. Moreover, an estimated negative mean market elasticity in an SMSA does not imply that monopsony does not characterize some part of that SMSA's labor market even though a negatively-sloped supply curve to a firm is inconsistent with that firm's being a monopsonist. The size of the estimated aggregate elasticity merely indicates what segment of a backward-bending supply curve the average individual in the SMSA is in. It is possible therefore for a significant proportion of the population to be on the positively-sloped part of their supply curves and face monopsonistic situations even though on average, the elasticity is negative. Whatever the sign of the mean SMSA elasticity, therefore, if there is a significant degree of monopsony in that market, there may still be a positive relationship between the mean elasticity and the average wage.

Equations (1) and (2) correspond to the estimating equations in the two stages: in equation (1), the labor-force participation variable L_{ijk}

$$L_{ijk} = \alpha_{ik} + \sum_{h=1}^{T} \beta_{ikh} X_{ijkh} + \beta_{ik\eta} \hat{W}_{ijk} + u_{ijk} \tag{1}$$

$j = 1 \ldots n_i$, $i = 1 \ldots S$, k = male, female

$$\ln\bar{W}_{ik} = a_k + \sum_{l=1}^{t} b_{kl}\bar{Y}_{ikl} + b_{k\eta}\hat{\beta}_{ik\eta} \frac{\bar{W}_{ik}}{\bar{L}_{ik}} + e_{ik} \tag{2}$$

of the j^{th} individual of sex k in the i^{th} labor market is regressed against that individual's vector of personal characteristics X_{ijkh} and predicted wage $\hat{W}_{ijk}$

The set of estimated mean elasticity estimates $\hat{\beta}\frac{\bar{W}_{ik}}{\bar{L}_{ik}}$ along with $\bar{Y}_{ik}$, the set of average sex-specific wage-determining characteristics for each labor market i, are then used as regressors in wage function (2) which is run across the S markets. The test for the existence of monopsony in the male and female labor markets is thus whether $b_{m\eta}$ and/or $b_{f\eta} > 0$.

a. The Sample

The data sets utilized, the Public Use Samples of the 1970 Census of Population, were chosen because they contain sufficient numbers of observations for the two-stage testing procedure and because they provide many of the relevant characteristics needed for the proper estimation of the supply and wage equations. Unfortunately, one important earnings determinant--labor-force experience--is not available from these tapes so that it was necessary to exclude all women who have ever been married from our samples. This was done because the usual computed proxy for work experience, age less years of schooling minus 5, is not an accurate one for these women and its use would therefore contaminate our coefficient estimates in the female labor-force participation and wage equations. Malkiel and Malkiel (1973) and Sawhill (1973) have shown that the computed work experience variable, however, is a reasonable estimate of the actual work history of single women and males. Moreover, since a greater proportion of adult males and single women work than married women, estimates of the potential wage of men and single women not currently in the labor force, based on the earnings of working males and non-married women, will be subject to less of a "selectivity bias" than would be obtained in a sample of married women.[9] We need such estimates of potential earnings to obtain unconditional supply elasticities.

For these reasons, only never-married, non-farm, not self-employed white females not currently in school, aged 19-66 and living in SMSA's from the 1:100 Census tape were chosen for our sample of women. The male sample, however, consisted of all men without regard to marital status who met all the other criteria applied to the female group. To obtain a male sample that was of comparable size to that of the single women, the 1:1000 tape was used as a data base.

Both the male and female samples were grouped by SMSA; those SMSA's containing samples of men or women numbering less than 50 were excluded from the final data sets in order that all the intra-SMSA labor-force participation equations could be run over enough observations to obtain reasonably precise coefficient estimates. This procedure yielded a final sample consisting of 95 SMSA's and a total of 18408 males and 30155 never-married women.

b. Intra-SMSA Supply Functions

To obtain unconditional male and female labor-supply elasticities for each of the 95 SMSA's a two-stage least-squares regression technique was utilized.[10] First, exogenous predicted wage variables, representing the potential wages of all individuals in the sample, were derived from estimating equations in which the natural logarithim of the hourly wage rates of all men or all women reporting earnings in 1969 (96 percent of all men, 87 percent of the never-married women) were regressed against schooling level, age, age squared, and an occupational demand index for each sex group in each SMSA. This latter variable was computed by dividing the male-female employment ratio of the individual's occupation in the SMSA by the male-female ratio for that occupation in the aggregate U.S. population; i.e., the index for an individual in SMSA i employed in occupation q is $[m_i/f_i]q/[m/f]q$.

Since our hypothesis is that if monopsony exists, then the wage paid and the supply elasticity of the individual will be correlated, it is necessary to assume that the supply elasticities are randomly distributed in the population in order that unbiased estimates of wage offers can be obtained. If this assumption holds, that is, if the responsiveness of labor supply to wage changes is uncorrelated with such personal characteristics as age, experience, education, then any error in the first-stage wage-predicting equation resulting from the exclusion of the (individuals') supply elasticity variable should be impounded exclusively in the residual error term.

The second-stage (unconditional) labor supply equations were run separately on all men and all women in each SMSA sample regardless of labor-force status. The participation measure used, annual hours worked, was computed by multiplying the reported weeks worked during 1969 by the weekly hours worked in the Census week in 1970 and was set equal to zero for individuals not in the labor force.[11] Annual hours was selected as the dependent variable rather than weeks worked during the year because the use of the latter variable would have resulted in a downward bias in the estimated predicted wage coefficient (and the elasticity), given a positive correlation between weeks worked per year and hours worked in the week. Moreover, Fuchs (1967) has shown that the error component in the hours variable computed in the manner described tends to be quite small. Thus any spurious negative association between the computed annual hours and hourly wage variables due to the measurement error in the latter being inversely related to measurement error in the former, is likely to be negligible.

The independent variables appearing in both the male and female supply equations in addition to the predicted wage are own non-earned income, age, age squared, and educational level. For the married, spouse-present men other variables were included to capture the intra-family substitution and additional income and household productivity effects. These are the non-earned income, age, and educational level of the wife and the number of children less than 6 years of age in the household. In addition, a dummy for marital history (1 = divorced or separated, 0 otherwise) was added.

Table 1 displays the results of the male and female supply equations obtained from regressions run on the total sample . Since the labor-supply elasticities to be used in the second-stage were obtained from regressions pertaining to each of the 95 SMSA's these results are presented for illustrative purposes only. The signs and significance of the coefficients conform in general to the ususal theoretical and empirical labor-supply findings in the literature. In the whole-sample male equation, both the own and wife's non-earned income coefficients, which capture the pure income effect on labor supply, are negative and significant at the 1 percent level; of the 95 SMSA male equations, 72 of the male non-earned income coefficients displayed a negative sign and 60 wife's non-earned income coefficients were negative. The probability of obtaining these results from a sample population in which the coefficients are randomly distributed are less than .01 and .05 respectively. The whole-sample coefficient of the child variable is positive and significant (1 percent) and the variable had a positive coefficient in 56 of the 95 regressions ($p = .18$), indicating that the presence of children in the household increases the labor-force participation of males, as found by Smith (1971).

The estimated country-wide elasticity for the males is positive; however, in 62 of the SMSA equations the computed elasticity was negative. These results contrast with Finegan's (1962) well-known estimates of a negative hours-elasticity, obtained from aggregate interindustry and interoccupational cross-sections based on the 1940 and 1960 Censuses. As Feldstein (1968) has argued, however, Finegan's findings may have been plagued by simultaneous equations bias. Our results are more consistent with those of Feldstein, who found that of his eleven labor-supply regressions run on separate samples of mainly male workers only 4 had significant negative (conditional) wage coefficients. To test for the existence of a backward-bending supply curve, a quadratic predicted wage was entered in the whole-sample male equation (column 4). The negative significant coefficient on this term indicates that the male supply curve does bend back, but at a turning point at the tail end of the sample wage distribution.[13] In the whole-sample never-married female equation, the non-earned income variable coefficient also displays the correct sign; in 85 of the 95 SMSAs negative coefficients were obtained for this variable ($p < .00001$). The mean elasticity for the individual sample is positive, but in 51 SMSAs the computed elasticity is negative. Again, the addition of the squared predicted wage term in the second column provides evidence of a backward-bending never-married female supply curve. As in the male group, however, the turning-point occurs at a wage at the high end of the observations in the sample.

As one means of checking the sensitivity of the elasticity estimates to the possibility of selectivity bias, conditional female supply elasticities were estimated from regressions run only on employed women. It was expected that the wage elasticity obtained from the sample of working women would be less than the unconditional elasticity estimate of .32 given that the participation decision is uninfluenced by the (negative) income effect. The computed conditional elasticity was .16, indicating that any downward bias in the unconditional estimate is not so severe as to make it fall short of this lower bound.

c. Inter-SMSA Earnings functions

The functional form of the earnings equations to which the computed SMSA elasticities are added is that formulated and applied by Mincer (1974) and used by Malkiel and Malkiel (1973) and Thomas Johnson (1969) and is given by (3):

$$\bar{Y}_{ik} = e^{a_k + b_{1k} S_{ik} + b_{2k} E_{ik} + b_{3k} \bar{E}^2_{ik} + b_{\eta k} \bar{\eta}_{ik}} (\bar{H}_{ik}) \qquad k = m,f \qquad (3)$$

$\bar{Y}_{ik}$, $\bar{S}_{ik}$, $\bar{E}_{ik}$, and $\bar{H}_{ik}$ are the mean levels of annual earnings, schooling, labor-force experience (= age - S - 5), and annual hours worked of all those employed in 1969 for the two sex groups in each of the 95 SMSAs. $\bar{\eta}_{ik}$ is the mean SMSA elasticity for each sex computed from the estimated labor supply equations. This latter variable, crucial for our test, is probably measured with considerable error so that a bias toward zero in the elasticity coefficient is expected. The direction of the biases produced in the other coefficients due to the errors in the elasticity variable can be computed according to the technique of Levi (1973) and are indicated below.

Because it was suspected that the unexplained variance in average earnings might vary inversely with SMSA size the Goldfeld-Quandt test was applied to the unweighted samples. It was concluded that both the male and female equations were characterized by this form of heteroscedasticity (5 percent level) and the observations were thus weighted by (SMSA size / $10{,}000)^{1/2}$.

Table 2 presents the results from the weighted OLS regressions on the 95 SMSA's. In the 'narrow' form of the earnings function for both men and women (specification I), the coefficients conform generally to those obtained by others; however, the rates of return to schooling (the b_{1k}) for both men and never-married women are higher than those that have been reported elsewhere. To ascertain if these results are a function of the use of the 1970 population, a data base not exploited as of yet in the literature, or are the product of the selection procedure limiting the sample to the larger SMSA's, the earnings equations were run across (male, female) individuals of similar characteristics from the 1:10,000 tapes but without regard to residence. The returns to schooling obtained were 7.8 % for males and 12.3% for the never married women.[14] It appears that the estimated male return to schooling is increased more significantly by the exclusion of individuals residing outside of the larger SMSA's than is the case for the never-married women.

Another quantitative difference in the earnings function results is that neither of the experience terms are statistically significant. However, together the two variables contribute significantly (5 percent level) to the explanatory power of the earnings equation for both sexes, indicating that the experience variables are highly collinear in the aggregate sample.

In specification II, the computed elasticity term is added to test for the existence of monopsony. As can be seen, the elasticity coefficient is positive and significant for the never-married women (5 percent level, one-tailed test) but does not attain significance for men. The addition of SMSA size (specification III), found by Fuchs (1967) to be an important determinant of inter-city wage differentials, does not alter this result. As the elasticity-earnings relationship does not appear to be significant for males, men are not subject to monopsonistic exploitation to the extent that women are. Thus because of the significant positive correlation between female supply elasticities and wage rates, it appears that the earnings of never-married women are in part lower than their male counterparts' due to the existence of monopsonistic discrimination.

Table 3 presents the direction of the measurement biases calculated for the women's specification III, according to Levi's (1973) technique. As expected, the coefficient on the elasticity variable is biased towards zero. The rate of return or education also has a downward bias, indicating that 13.1 percent is an underestimate of the true rate of return to education for never-married women.

Finally, Table 4 presents our estimates of the magnitude of the combined effect of taste plus monopsonistic discrimination in column four. Even using a sample of never-married women, and after adjusting for differences in endowments, that is for elasticity, schooling, experience, hours worked, and city size we find an unexplained earnings differential of about 35% in our sample. Our estimate is in the same range as Oaxaca (1973) [.29], Fuchs (1971) [.34], higher then Malkiel and Malkiel (1973) [.19] and lower than Sawhill (1973) [.43]. However, we have shown that this computed residual differential cannot be interpreted as a measure of taste discrimination only, but in part can be attributed to the existence of monopsony.

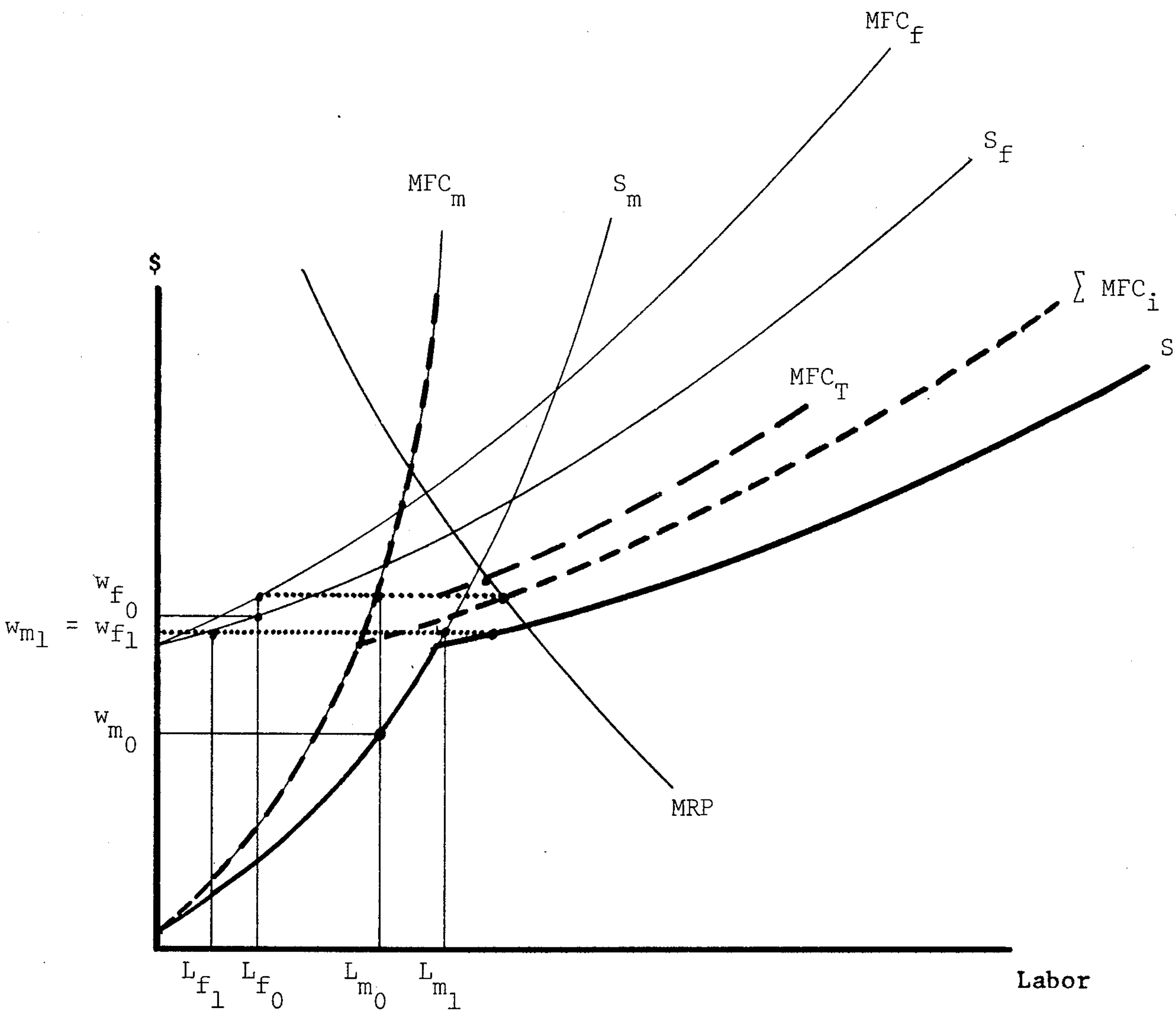
Table 5 shows that much of the differential in earnings is due to the lower return to education and experience of never-married women, as well as to women's fewer years of experience. The constant term of the regression, or the part of the wage differential which is not explained by the variables used, goes in favor of women.

Conclusion

We have shown in the theoretical section of our paper that the usual residual measures of the "taste" discrimination component of male-female wage differentials found in the literature were likely to be overestimates of this form of discrimination. "Taste" discrimination on the part of employers was shown to be one of two possible immobilizing pre-conditions for the monopsonistic exploitation of women, which could result in lower wages for women relative to men if the market for male labor were competitive. Since the immobility of women could also result from their socialized behavior, there is thus no way to evaluate the importance of taste versus monopsonistic discrimination in accounting for sex differences in earnings, controlling for personal characteristics.

Our empirical analysis, based on estimated supply curves in 95 labor markets, supports the hypothesis that some part of the unexplained sex differential in wages can be attributed to monopsonistic discrimination. Because we could not use measures of supply elasticities to individual firms, our analysis is thus a strong test for the existence of monopsony. Therefore, it must be noted that the positive and significant relation between the supply elasticities and earnings of never-married women does not indicate that all firms are monopsonists, but only that enough firms act as discriminating monopsonists with respect to these women so as to show up in our data. Moreover, it should be emphasized that our finding of no significant relationship between the market supply elasticities of males and male wages cannot be interpreted to mean that no males are subject to monopsonistic exploitation but only that we could not reject the null hypothesis of competitive male labor markets.

Finally, what does our conclusion concerning the importance of monopsonistic discrimination with respect to never-married women imply about the earnings of married females? While _a priori_ we might expect married women to be less geographically mobile than single women and thus the wage-supply elasticity relationship to be more significant in a statistical sense for this group, we can not say to what extent married women experience monopsonistic wage reductions relative to single women without testing for the existence of monopsonistic discrimination based on marital status as well as sex. Given this additional form of monopsonistic discrimination, our results suggest that married women might receive higher wage rates than single females, as a result of their relatively high elasticity of supply to the labor market.



MFC$_f$
S_f
MFC$_m$
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w_{f_0}
$w_{m_1} = w_{f_1}$
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MRP
L_{f_1}
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Labor

Figure 1

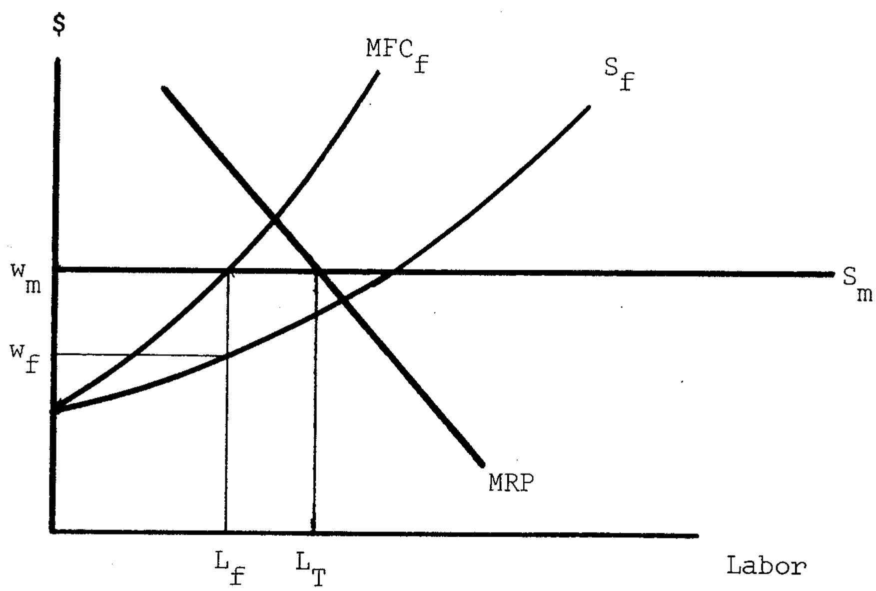

Figure 2

Table 1

Estimates of Unconditional Supply Equations (Annual Hours Worked)

Countrywide Sample

Variable	Coefficients and t-ratios Women		Men	
WAGE	32.9721 (15.800)**	103.9272 (12.543)**	124.8304 (14.068)**	368.0092 (10.218)**
$WAGE^2$		-.28022 (8.848)**		-3.0808 (6.959)**
EDUC	45.8811 (33.032)**	36.0870 (20.333)**	4.35680 (1.818)	-30.6303 (5.502)**
AGE	59.7063 (25.329)**	60.3683 (25.629)**	64.3480 (19.300)**	50.5419 (13.040)**
AGE^2	-.12696 (25.495)**	-.79143 (26.478)**	-.97902 (26.489)**	.99811 (26.961)**
NEINC	-.12696 (27.362)**	-.17243 (24.918)**	-.10317 (20.229)**	-.16532 (16.082)**
AGEW			3.1519 (4.356)**	3.1624 (4.376)**
EDUCW			13.8718 (7.129)**	13.7654 (7.081)**
NEINCW			-.00983 (2.209)**	-.01177 (2.645)**
STATUS			56.2013 (1.701)	58.2694 (1.760)
CHILD			31.5404 (3.789)**	30.7519 (3.698)**
C	-426.4518	-526.0909	-3.5727	-152.6279
R^2	.090	.092	.138	.140
n	18408	18408	30155	30155

** Significant at the 1 percent level (two-tailed test)

Source: 1970 Public Use Sample

Table 2

Estimates of Earnings Functions (ln of Annual Earnings)

Variable	Coefficients and t-ratios					
	Women			Men		
	I	II	III	I	II	III
EDUC	.12340 (5.950)**	.12642 (6.150)**	.13070 (6.744)**	.13762 (7.557)**	.13813 (7.466)**	.13897 (7.463)**
EXP	-.00082 (0.020)	-.00144 (0.035)	-.00016 (0.018)	.03877 (0.420)	.03892 (0.419)	.04518 (0.482)
EXP^2	.00111 (0.650)	.00113 (0.669)	.00083 (0.519)	-.00046 (0.231)	-.00046 (0.230)	-.00061 (0.302)
LNHRS	.40038 (2.885)**	.38810 (2.830)**	.48788 (3.690)**	.61538 (3.370)**	.61703 (3.360)**	.64304 (3.393)**
η		.03613 (1.771)*	.03177 (1.652)*		-.00295 (0.196)	-.00320 (0.212)
SMSA			.00013 (3.250)**			.00002 (0.500)
C	3.71640 (3.664)**	3.75628 (3.700)**	2.88426 (2.960)**	2.07390 (1.090)	2.05248 (1.068)	1.76079 (0.885)
R^2	.443	.453	.569	.492	.496	.542

** Significant at the .5 percent level (one-tailed test)

* Significant at the 5 percent level (one-tailed test)

Source: 1970 Public Use Sample

Table 3

Bias in Coefficients of Women's Equation Due to Errors in Measurement of Elasticity[1]

Variable	EDUC	EXP	EXP^2	LNHRS	η	SMSA	C
Bias	-	+	-	+	-	+	-

[1]See Levi (1973)

Table 4

Predicted Annual Earnings, Based on Earnings Function[1]

Specification		Characteristics of women	Characteristics of men	Ratio	Monopsony, Taste Discrimination Earnings Differential
		(a)	(b)	(a)/(b)	1-(a)/(b)
Women's Function III	(e)	$4912	$ 6863	.715	.28
Men's function III	(f)	$7726	$ 10541	.734	.27
Ratio (e)/(f)		.636	.651		

[1]Across 95 SMSA's

Table 5

Component Differentials

Variable	$e^{X_f(\beta_f - \beta_m)}$	$e^{\beta_m(X_f - X_m)}$
EDUC	.8864	1.1077
EXP	.5741	.6155
EXP^2	1.2488	1.2595
n	.9989	.9998
SMSA	1.0131	1.0
C	3.0756	0

$$W_f/W_m = e^{\sum_0^6 X_f(\beta_f - \beta_m)} \, e^{\sum_0^6 \beta_m(X_f - X_m)} \, HRS_f^{.4879-.6430} \left(\frac{HRS_f}{HRS_m}\right)^{.6430}$$

$$= e^{\sum_0^5 X_f(\beta_f - \beta_m)} \, HRS_f^{-.1551} \, e^{C_f - C_m} \, e^{\sum_0^6 \beta_m(X_f - X_m)} \cdot \left(\frac{HRS_f}{HRS_m}\right)^{.6430} = D \cdot U \cdot E$$

where D is the effect of the differences in the regression coefficients,

U is the unexplained term (the constant) in the regression equation,

E is the effect of differences in endowments.

D = .207, U = 3.076, E = .733, D.U. =.636

Footnotes

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[1]p. 697-698.

[2]p. 150.

[3]See, for instance, Madden, forthcoming 1975.

[4]We ignore the game theoretical problems of dealing explicitly with degrees of monopsony power. In the theory, we look at absolute monopsony power and none. In the empirical section, we assume that varying degrees of monopsony power affect wages in a continuous manner.

[5]The effect of pure taste discrimination can be measured by the differential when all markets are perfectly competitive,in wages paid to two workers who equally productive (e.g., the same education, experience etc.). See Oaxaca (1973) and Sawhill (1973) for examples of attempts to measure the effect of discrimination by controlling for personal characteristics. They confound, however, the effects of taste and monopsonistic discrimination.

[6]We are considering only the case where the discriminating monopsonist employs both types of labor.

[7]There is one complication: A wage equalization policy might result in segregation, whereas the discriminating monopsonist hired from each group. The wage of the group employed will not fall below its wage paid by the discriminating monopsonist.

[8]Samuelson discusses the analogous case of a discriminating and simple monopolist, Foundations of Economic Analysis, pp. 42-45.

[9]For a discussion of the possible biases resulting from the estimation of the potential wage of non-working women from a sample of employed women see

Gronau (1974). Heckman (1974) proposes a maximum-likelihood iterative method which yields consistent and asymptotically unbiased parameter estimates for computing the potential wage offers of non-employed individuals. Because the small sample properties of such a technique are not known, however, given the size and number of our samples and the exclusion of all ever-married women, we did not attempt this procedure.

[10]The reason for estimating a predicted wage is to obtain unconditional supply elasticities rather than to eliminate a simultaneity problem. With micro data the dependence of the wage received on the individual's quantity of work is negligible. Thus the wage-producing equation is not derived from a formal simultaneous equation's system.

[11]Thus the dependent variable is truncated normal. Amemiya's (1974) iterative maximum likelihood estimation technique, which he demonstrates yields strongly consistant and asymptotically normal parameter estimates when the dependent variable displays this property, was not applied. See note 9.

[12]Supply equations using annual weeks as the dependent variable were estimated for a random sub-sample of the SMSA's. The coefficients were generally significantly less precise in these equations than in the comparable hours regressions.

[13]Bognanno et. al. (1974) obtained similar results with a quadratic wage term in their estimates of the conditional supply elasticities of married female nurses.

[14]Mincer and Polachek (1974) using the 1967 National Longitudinal Survey estimated a rate of return to schooling of non-married women of 7.7%

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