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PAY DIFFERENCES BETWEEN WOMEN'S AND MEN'S JOBS: THE EMPIRICAL FOUNDATIONS OF COMPARABLE WORTH LEGISLATION

George Johnson

Gary Solon

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ABSTRACT

Civil rights legislation of the 1960s made it illegal for an employer to pay men and women on different bases for the same work or to discriminate against women in hiring, job assignment, or promotion. Two decades later, however, the ratio of women's to men's earnings has shown little upward movement. Furthermore, major sex differences in occupational distribution persist with predominantly female jobs typically paying less than predominantly male jobs. This negative relationship between wage rates and femaleness of occupation has stimulated efforts, in both the judicial and political arenas, to establish "comparable worth" procedures for setting wage rates.

This paper estimates the relationship between wages and femaleness of occupation and finds that it is indeed negative even after controlling for relevant worker and job characteristics. The magnitude of the relationship, however, implies a surprisingly small effect for a comprehensive comparable worth policy. The estimates indicate that, even if comparable worth succeeded in eliminating this negative relationship, the disparity between mean male and female wages would be reduced by well under ten percent of its current magnitude.

George Johnson
Department of Economics
The University of Michigan
Ann Arbor, MI 48109
(313) 763-4390

Gary Solon
Department of Economics
The University of Michigan
Ann Arbor, MI 48109
(313) 763-1306

PAY DIFFERENCES BETWEEN WOMEN'S AND MEN'S JOBS: THE EMPIRICAL FOUNDATIONS OF COMPARABLE WORTH LEGISLATION

Civil rights legislation of the early 1960s made it illegal for an employer to pay men and women on different bases for the same work or to discriminate against women in hiring, job assignment, or promotion. Two decades later, however, the ratio of women's to men's earnings has shown little upward movement. Furthermore, major sex differences in occupational distribution persist with predominantly female jobs typically paying less than predominantly male jobs. For example, in 1983 the percentages female among truckdrivers and secretaries were respectively 3.2 and 99.0, with the truckdrivers receiving considerably higher wages. Similar comparisons can be made of engineers versus librarians, professors of economics versus professors of art history, and literally hundreds of other examples of primarily men's jobs (MJs) versus primarily women's jobs (WJs).

Such comparisons have generated much discussion, in both the judicial and political arenas, concerning the desirability of establishing "comparable worth" (CW) procedures for setting wage rates.[1] The presumption underlying the CW movement is that the observed negative relationship between wage rates and femaleness of an occupation reflects an "undervaluation" by society of WJs relative to MJs. Accordingly, CW legislation -- either partial (applied to a particular branch of government and/or a subset of private employers) or comprehensive -- would require that an employer's wage structure across

[1]See Bureau of National Affairs (1981) for a useful review of the comparable worth doctrine and its relation to existing legislation, particularly the Equal Pay Act and Title VII of the Civil Rights Act. Recent progress by the comparable worth movement is summarized in Goodman (1984).

jobs be justified by "job evaluation" procedures. Specifically, each job within an organization would be assigned points in each of several dimensions (such as skill requirements, responsibility, effort, and working conditions), and these scores would somehow be aggregated to an overall index for determining relative compensation levels.[1] Under the presumption that WJs are currently undervalued, a legal requirement that relative wage levels be set in this manner, rather than by the present combination of market and institutional forces, would tend to reduce the disparity in the remuneration of MJs and WJs.

The primary object of this paper is to estimate and analyze the negative relationship between wage rates and femaleness of an occupation. This relationship, which is the empirical basis for CW legislation, is further discussed and couched in a convenient econometric framework in Section I. Section II presents an empirical analysis of the relationship, and Section III explores how the relationship differs between the public and private sectors. Section IV summarizes the implications of the results for CW policy.

I. Conceptual Issues

The presumed stylized fact that motivates CW proposals is the negative relation, ceteris paribus, between an individual's hourly earnings (W, which we express logarithmically) and the fraction of workers in his or her occupation who are women (F). Specifically, we posit that

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[1]Of course, this job evaluation process is extremely problematical. A detailed discussion in Treiman and Hartmann (1981) notes that "there are no definitive tests of the 'fairness' of the choice of compensable factors and the relative weights given to them," and concludes that the process is "inherently judgmental."

(1) $W = f(Z,F,S)$

where Z is a vector of characteristics that influence earnings (human capital, location, working conditions, etc.) and S denotes the individual's sex. This earnings function is depicted in Figure 1. The vertical axis measures the average log wage rates for men and women given their average values of the Z's. These wages $W_s|\bar{Z}_s$ are assumed to be negatively related to F on the horizontal axis. The figure as drawn assumes that, given F, men typically earn more than women. This would occur if men have higher-earning average characteristics or if women are subject to wage discrimination.

In the framework of Figure 1, one can view the relevant civil rights provisions of the 1960s as an attempt to move the men's and women's functions, net of average differences in the Z's, into confluence. Realization of that goal, however, would still leave women with lower pay than men because women would still tend to be concentrated in occupations with larger F. The CW proposals of the 1980s can be viewed as an attempt to remedy this situation by making $\partial W/\partial F=0$, that is, by making Figure 1's earnings functions horizontal.

Although the sources of the functions' negative slopes are not the main focus of this paper, they warrant some discussion. The two most prominent explanations for the negative relationship between wages and femaleness of occupation are sex differences in preferences and occupational exclusion.[3] According to the preferences explanation, workers have heterogeneous tastes for job chararacteristics, and the

[3]A third explanation is that wages for some WJs, especially hospital nurses, are depressed because employers have monopsony power. There is little evidence, though, that this analysis applies to enough WJs to provide a general explanation.

distribution of these tastes differs between men and women. These gender differences in job preferences might arise from socialization concerning sex roles and "appropriate" work (the nurse/doctor syndrome), differing family responsibilities, or differing expectations regarding continuous versus intermittent labor market attachment. In any case, such taste differences would lead to sex differences in occupational distribution. Furthermore, if the preference distributions were such that the supply of workers to female-dominated occupations were especially large relative to the demand, these occupations would pay relatively low wages. In this analysis, the negative relationship between wages and femaleness of occupation results from voluntary choices, and the wage structure cannot be "improved" by legislative fiat. The relative wage changes mandated by a CW policy would be undesirable on efficiency grounds and not especially appealing on equity grounds.[4]

An alternative explanation for $\partial W/\partial F<0$ is that women are systematically excluded from many occupations and are thus "crowded" into a subset of occupations with depressed wage rates.[5] According to this explanation, women become secretaries, librarians, and art historians not by choice, but because they are blocked from becoming truckdrivers, engineers, and economists. The employer practices that would lead to such occupational exclusion are illegal under Title VII of the Civil Rights Act, and vigorous enforcement of that statute appears to be a more suitable long-run remedy than CW legislation. On the other hand, even the complete elimination of occupational exclusion would

[4]A more formal analysis is presented by Killingsworth (1984).

[5]See Bergmann (1971, 1974).

mainly benefit new female workers and might be of little help to older women already committed to traditionally female career paths. CW policy might then be advocated on equity grounds as a means of compensating the latter group.

Whatever the <u>source</u> of the negative relationship between wages and femaleness of occupation, the remainder of this paper attempts to <u>quantify</u> that relationship and thereby to ascertain the potential impact of CW on women's wages relative to men's. We estimate for each sex the following linearized version of equation (1),

$$(2) \qquad W = \beta_S' Z + \gamma_S F + \epsilon,$$

where β_S is a vector of parameters (for sex S) associated with the characteristics vector Z and ϵ is a disturbance term. For an individual of sex S, $\gamma_S = \partial W / \partial F$ and exp (γ_S) represents the <u>ceteris paribus</u> wage ratio between a virtually all-female occupation and a virtually all-male occupation.[6] In the present formulation, then, the CW goal of making $\partial W/\partial F = 0$ can be described as setting $\gamma_S = 0$.

The average log wage for each sex is $\bar{W}_S = \beta_S' \bar{Z}_S + \gamma_S \bar{F}_S$, and the male-female difference in average log wages is

$$(3) \qquad D = (\beta_m' \bar{Z}_m - \beta_f' \bar{Z}_f) + (\gamma_m \bar{F}_m - \gamma_f \bar{F}_f).$$

Since the goal of CW is to eliminate the negative partial relationship between wages and femaleness of occupation, we can obtain an upper bound on CW's relative wage impact by determining how much D would be reduced if the second term in equation (3) were eliminated by setting $\gamma_m = \gamma_f = 0$.

[6] Clearly, our empirical specification is influenced by Oaxaca's (1973) seminal paper.

(The reasons why even comprehensive CW legislation is unlikely to reach this upper bound will be discussed below.) The next section describes our efforts to estimate the γ's and CW's potential impact.

II. Empirical Analysis

To estimate equation (2), we use May 1978 Current Population Survey (CPS) data on nonagricultural wage and salary workers, at least 16 years old, who responded to the supplementary questions concerning their "usual weekly earnings" and "usual weekly hours." Relative to other surveys, the CPS has the advantage of detailed occupational information on a large national sample. Relative to subsequent CPS samples that were asked the supplementary questions, the May 1978 sample has the advantages of a larger size and a more "normal" unemployment situation.

The dependent variable in our regression analyses is the natural logarithm of the ratio of usual weekly earnings to usual weekly hours. This variable, W, averages 1.757 for the 24,056 men in our sample and 1.346 for the 19,412 women. This implies that D in equation (3) is 1.757-1.346=.411, which means that the (geometric) mean wage for women is 33.7 percent less than that for men. Our goal is to estimate how much of this wage difference is attributable to femaleness of occupation and how much to other factors.

Our femaleness variable, F, is the proportion female in the individual's three-digit occupation. Whenever possible, these values were obtained from 1978 CPS data as reported in the January 1979 Employment and Earnings. In occupations for which the CPS cell sizes were too small to allow publication, we had to extrapolate from 1970 census data. The resulting F variable ranges from less than .01 (for such categories as firefighters, automobile mechanics, and plumbers and pipefitters) to over .99 (secretaries and housekeepers).

The other variables, comprising the Z vector in equation (2), are: years of school completed; potential work experience (age minus years of school minus six) and its square; regional dummy variables for Northeast, North Central, and West; three dummies for residence in a large, medium, or small SMSA; race dummies for black or other minority race; dummies for voluntary and involuntary part-time work; two marital status dummies, one for married and the other for separated, widowed, or divorced; number of children and a dummy for presence of children; dummies for union membership and, if not a member, coverage by a union contract; a government employment dummy; dummies for 20 major industries (mining, construction, durable goods manufacturing, etc., with private households as the omitted category); the fraction of workers in the individual's occupation who worked less than 30 hours in the 1970 census week; and a series of occupational characteristics indices, developed by the staff of a National Research Council committee,[7] describing the "general educational development," "specific vocational preparation," strength, physical demands, and undesirable environmental conditions associated with the occupation. The committee itself acknowledged that these latter variables are arbitrary and of dubious validity; on the other hand, they are qualitatively similar to the job evaluation scores that would be applied under a CW policy and therefore ought to be controlled for in the estimation of the γ's.

Indeed, the general question of whether a variable should be included in the Z vector comes down to this issue of whether CW would

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[7]Miller et al. (1980, appendix F).

allow an employer to base pay on that variable. For example, the industry dummies undoubtedly belong in the regressions because CW would require equal pay for work of comparable worth only <u>within</u> firms. Wage differences <u>between</u> firms would not be covered, and <u>a fortiori</u> industry wage differences would be unaffected. On the other hand, whether union status belongs in Z depends on whether CW would permit an employer to pay union workers in one occupation on a different basis than nonunion workers in another occupation. Presumably, variables such as race and marital status do not belong in the regressions except as proxies for other characteristics that employers would be allowed to use. Given the considerable ambiguity in the choice of Z variables, we will make careful efforts in the analysis below to clarify how different control variables affect the estimation of the γ's.

We begin by estimating, for each sex, simple regressions of W on F. The results, which describe the <u>gross</u> relationship between wages and femaleness of occupation, are reported in columns 2 and 5 of Table 1. For men, the estimated coefficient for F, $\tilde{\gamma}_m$, is -.343, which implies a .710 wage ratio between virtually all-female and all-male occupations. For women, the estimated coefficient $\tilde{\gamma}_f$ is -.244, implying a .783 wage ratio. These results confirm Treiman and Hartman's (1981) findings from aggregate data that earnings are negatively related to femaleness of occupation for both men and women and that the relationship is stronger among men.

These estimates, however, take no account of differences in workers' characteristics or the characteristics of their jobs. Columns 3 and 6 in Table 1 report the results of estimating multiple regressions of W on F controlling for the full Z vector described above. In general, the estimated coefficients of the control variables are unremarkable. It is worth mentioning, though, that the coefficient estimates for schooling and experience appear smaller than most in the earnings function literature simply because the wage effects of schooling and experience are partly absorbed by the coefficients of the "general educational development" and "specific vocational preparation" occupational variables.*

What is more striking is that the inclusion of the Z variables reduces the estimated F coefficients by more than half. For men, controlling for Z reduces the estimated coefficient of F from $\tilde{\gamma}_m$=-.343 to $\hat{\gamma}_m$=-.168; for women, the reduction is from $\tilde{\gamma}_f$=-.244 to $\hat{\gamma}_f$=-.090. The new coefficient estimate for men implies a .845 ceteris paribus wage ratio between virtually all-female and all-male occupations. The new coefficient estimate for women implies a ratio of .914.

Two important questions arise concerning the interpretation of these coefficient estimates. First, if we take the multiple regression estimates at face value, what do they imply about the potential impact of CW? As discussed above, CW can be viewed as an attempt to set $\gamma_m=\gamma_f=0$ and thereby to eliminate the second term in equation (3) for D, the male-female difference in average log wages. A simple computation using $\hat{\gamma}_m$, $\hat{\gamma}_f$, and the sample means of F_m and F_f estimates this term to be .029, as compared to a total D of .411. Given these estimates of the γ's, then, even total elimination of $\partial W/\partial F$ would reduce D by only about 7 percent. To put it another way, whereas the (geometric) mean wage was

*The "wrong" sign of some of the other occupational characteristics' coefficient estimates is a common result (see Brown (1980)).

previously 33.7 percent less for women than for men, eliminating $\partial W/\partial F$ would merely change this figure to 31.8 percent.

Of course, all this rests on our multiple regression estimates $\hat{\gamma}_m$ and $\hat{\gamma}_f$. If instead we used the simple regression estimates $\tilde{\gamma}_m$ and $\tilde{\gamma}_f$, we would estimate the second term of D to be .102. Even then, setting $\gamma_m=\gamma_s=0$ would leave most of D remaining, but the difference from the multiple regression implications is considerable. This raises the second question of how the inclusion of various control variables affects the estimation of the γ's. To clarify the influences of different variables, we use the fact (demonstrated in the appendix) that the change from a simple to a multiple regression estimate of γ can be expressed as

$$(4) \qquad \hat{\gamma} - \tilde{\gamma} = - \sum_{j=1}^{K} \hat{\beta}_j \, b_{jF}$$

where the $\hat{\beta}_j$'s are the estimated coefficients of the K control variables in the log wage regression and the b_{jF}'s are the coefficients from auxiliary simple regressions of the control variables on F. Equation (4) enables a straightforward decomposition of the difference $\hat{\gamma}-\tilde{\gamma}$ into the parts attributable to each control variable.

Table 2 summarizes the decompositions for $\hat{\gamma}_m-\tilde{\gamma}_m$ and $\hat{\gamma}_f-\tilde{\gamma}_f$. The striking finding is that, for both men and women, the bulk of the reduction from $\tilde{\gamma}$ to $\hat{\gamma}$ is due to the industry dummy variables. A closer look at the underlying data reveals the main reason for this industry effect: workers in construction and manufacturing (especially durables) are relatively well-paid given their other characteristics and are predominantly male. Once the wage effect of belonging to these industries is separately accounted for, the remaining effect of femaleness of occupation is considerably reduced.

It is crucial to understand that, even if the relation between wages and femaleness of <u>industry</u> arises from discriminatory exclusion practices, CW would not remedy the resulting pay differences. As discussed above, CW would require equal pay for work of comparable value only within firms. Interindustry wage differences would be unaffected. Consequently, the smaller γ estimates from the multiple regression analyses are clearly better indicators of the potential impact of CW. Indeed, if anything, they probably <u>exaggerate</u> CW's impact because they control only for industry effects and not firm effects. An ideal data base that enabled controls for firm effects as well could also take account of CW's inability to alter pay differences between firms in the same industry.[9]

Although of much less importance than the industry variables, the dummy variables for union membership and coverage account for about 15 percent of $\hat{\gamma}-\tilde{\gamma}$. This reflects the facts that there is a large estimated union/nonunion wage differential for both men and women and that union organization is much more prevalent in MJs than in WJs. It is also interesting to note that the portion of the average log wage differential between men and women that can be attributed to these union variables is .30 x .209 + .03 x .113 - .15 x .214 - .04 x .128 = .028,

[9]Treiman and Hartmann (1981, pp. 39-40) summarize several studies' findings that, within occupations and industries, women tend to be concentrated in lower-paying firms. CW would not address this source of wage differences. Indeed, if CW legislation were enacted, its inapplicability to interfirm differences might be exploited through business reorganizations. For example, a firm might "contract out" its female-dominated clerical functions to another firm to preclude pay comparisons with its other job categories.

nearly the same portion that we estimated as CW's potential impact. In other words, a policy that (somehow) eliminated the relative wage impact of union membership and coverage would have about the same effect on the female/male wage ratio as would economy-wide CW legislation.

It could certainly be argued that some of the variables included in Table 1's multiple regressions should in fact be excluded. For example, a stringent CW law might require firms to pay their nonunionized employees comparably with their unionized employees. If so, the unionization variables should be excluded from the regressions. Similarly, although some variables such as those for marital status may proxy for legitimate determinants of pay, they may not belong in the regressions in their own right.

Therefore, to check the robustness of our estimates of $\partial W/\partial F$, we have also estimated a parsimonious model that controls only for schooling, experience, region, SMSA, government employment, industry, and occupational characteristics other than fraction part-time. The results are very similar to those in the full model. For men, the estimated γ is -.176 (.015), as compared to -.168 in the full model. For women, the estimated γ is -.085 (.014), as compared to -.090. Eliminating these γ's through CW would reduce D, the male-female difference in average log wages, by $.085 \times .71 - .176 \times .21 = .023$, even less than the .029 estimated in the full model.

The most important implication of these results is that, since CW would not apply across industries (or, indeed, across firms), it is unlikely to eliminate a major fraction of the disparity between women's and men's wages. If the model is estimated without the industry dummy variables, the estimated γ's are somewhat greater, implying greater effects from CW.[10] As far as we know, however, interindustry wage equalization is beyond the scope of any proposed or imagined CW legislation.

III. Differences between the Public and Private Sectors

Most of the legislative and judicial activity with respect to CW has been in the public sector. Several states and municipal governments have opted or been forced by the courts to adopt CW procedures for determining pay scales across occupations, and a current bill in Congress calls for the same in the federal civil service. Legislation applying CW to the entire economy (i.e., with the same coverage as the Equal Pay Act and Title VII of the Civil Rights Act) seems at least several years from enactment. It is, therefore, interesting to see if the negative ceteris paribus relation between wage rates and femaleness of jobs is stronger in the public than in the private sector. Is there an empirical justification for the current concentration of CW activity in the public sector?

Table 3 reports the estimated γ's when the CPS sample is divided into public and private employees. The simple and multiple regression results in the first two rows are analogous to the full-sample results in Table 1.[11] For both men and women, the multiple regression estimates

[10]In fact, the estimates of γ in the full model exclusive of industry dummy variables are -.210 (.015) for men and -.105 (.013) for women, not drastically larger than those that do account for industry. It turns out that the occupational characteristics variables are rather highly correlated with industry and thus account for a large portion of the reduction in $\hat{\gamma}$ when the industry variables are omitted.

[11]One surprising result in Table 3 is that, for male public employees, the simple regression estimate of γ is much smaller in magnitude than the multiple regression estimate. Equation (4)'s method for assessing the influence of various control variables reveals large positive values of $\hat{\beta}_j b_{jF}$ for years of school and the occupational

of the γ's are indeed larger in the public than the private sector, although the differences are not statistically significant at conventional levels. These γ estimates imply that a CW-induced elimination of $\partial W/\partial F$ would reduce D, the average male-female log wage differential, by .042 in the public sector and .037 in the private sector. Both of these figures are larger than our full-sample estimate of .029, but they are still small relative to the public sector D of .293 and the private sector D of .451.

In the previous section, we argued that estimation of γ's should control for industry effects, but this is not entirely clear in the public sector. A state government, for example, <u>might</u> declare that its employees would be paid "comparably" across department lines -- in public administration, hospitals, transportation, education, etc. To explore this possibility, the third row of Table 3 reports the public sector γ's estimated without controlling for industry. These estimates are considerably higher than those with industry controls, and they imply that eliminating $\partial W/\partial F$ in the public sector would reduce that sector's D by .064. On the other hand, all our public sector estimates may exaggerate the impact of CW because they do not control for governmental unit. Even if a state government, for instance, implemented CW across department lines, it would not have to pay comparably to the federal government, municipalities, or other state governments. Just as the previous section's estimates overlook CW's

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characteristics indices. In other words, male government employees with more education and higher-earning occupational characteristics tend to work in occupations with relatively high fractions of women. Thus, although the simple relation between W and F is small for this group, the relation becomes much larger when schooling and occupational characteristics are controlled for.

inability to affect wage differences between firms, this section's public sector estimates overlook CW's inapplicability to intergovernmental wage differences.

IV. Conclusions

This study has measured the relationship between wages and femaleness of occupation. When relevant worker and job characteristics, including industry effects, are controlled for, a negative relationship between wages and femaleness of occupation still appears, in accordance with the concerns of the comparable worth movement. The magnitude of the relationship, however, implies a surprisingly small effect for a comprehensive CW policy. Our main estimates indicate that, even if CW succeeded in eliminating this negative relationship, the disparity between mean male and female wages would be reduced by well under ten percent of its current size, and we believe that, if anything, these estimates overstate CW's impact. These findings may disappoint CW advocates who expect CW to achieve drastic changes in the U.S. relative wage structure; correspondingly, they may soothe the fears of CW opponents who view it as the worst idea since minimum wage legislation.

At various points in the paper, we have presented statistical calculations of CW's potential impact on the male-female wage differential. These are useful for assessing CW's initial effects, but they should not be taken as predictions of its ultimate impact on the relative earnings of men and women. The most plausible model of CW's long-run effects is one in which only a fraction of WJs are covered by the law. (At present, CW activities are limited to a portion of the public sector; even a comprehensive CW law would, like existing fair employment legislation, be effectively confined to the public sector plus large private firms.) As Killingsworth (1984) and Ehrenberg and Smith (1984) have noted, given downward-sloping demand curves, any policy that raises the wages of covered WJs relative to MJs and uncovered WJs will lower employment in the former relative to the latter. Thus, CW would raise the wages of covered WJs, lower the wages of uncovered WJs, and increase the fraction of all WJs in the low-wage uncovered sector.[12] Whether women in the aggregate would gain or lose,, either in absolute terms or relative to men, depends on the size of the relevant demand elasticities.[13]

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[12]This result is similar to those concerning the impact of unionism and minimum wage legislation in Johnson and Mieszkowski (1970) and Welch (1974).

[13]Recent estimates of multi-factor partial elasticities of complementarity, as in Grant and Hamermesh (1981), suggest that the long-run demand elasticities may be rather high, in which case women would lose from CW. In addition, unless a comprehensive CW law contained strict provisos against "contracting out" (as described in footnote 9), the reduction in WJs in the covered sector might be very large indeed.

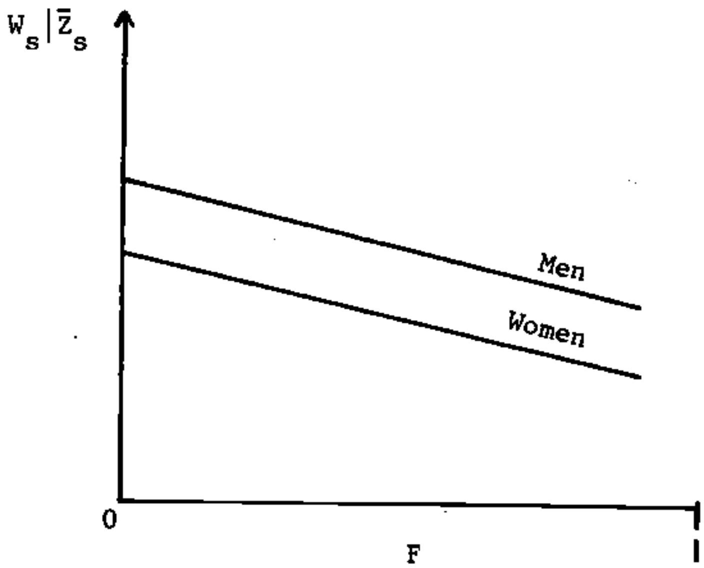

Figure 1

Table 1. Estimated Coefficients (and Standard Errors) in Log Wage Equations

Variables	Men			Women		
	Means	Regression Coefficients		Means	Regression Coefficients	
		Simple	Multiple		Simple	Multiple
Intercept	1.0	1.829 (.005)	.372 (.054)	1.0	1.520 (.010)	-.175 (.040)
F	.21	-.343 (.015)	-.168 (.015)	.71	-.244 (.013)	-.090 (.014)
Schooling	12.4		.046 (.001)	12.3		.039 (.001)
Experience (potential)	18.0		.023 (.001)	17.4		.012 (.001)
$Experience^2$	527.6		-.0004 (.00001)	513.7		-.0002 (.00002)
Northeast	.21		.021 (.008)	.21		.012 (.008)
North Central	.27		.048 (.007)	.27		-.001 (.008)
West	.23		.113 (.008)	.23		.076 (.008)
Large SMSA	.13		.124 (.009)	.13		.170 (.009)
Medium SMSA	.21		.118 (.007)	.21		.104 (.007)
Small SMSA	.25		.037 (.007)	.25		.038 (.007)
Black	.07		-.087 (.010)	.10		.018 (.010)
Other minority	.02		-.020 (.018)	.02		.063 (.019)

Table 1 (continued)

Variables	Men			Women		
	Means	Regression Coefficients		Means	Regression Coefficients	
		Simple	Multiple		Simple	Multiple
Voluntary part-time	.08		-.100 (.012)	.24		-.049 (.007)
Involuntary part-time	.02		-.100 (.017)	.05		-.058 (.013)
Married	.71		.120 (.008)	.56		.042 (.008)
Separated, widowed, divorced	.07		.078 (.012)	.19		.051 (.010)
Children dummy	.55		.018 (.009)	.51		-.022 (.009)
Number of children	1.13		-.001 (.003)	1.01		-.011 (.003)
Union member	.30		.209 (.006)	.15		.214 (.008)
Union coverage	.03		.113 (.016)	.04		.128 (.015)
Government	.18		-.069 (.012)	.23		.058 (.011)
Mining	.02		.500 (.047)	.003		.685 (.057)
Construction	.10		.381 (.044)	.01		.439 (.034)
Manuf., durables	.19		.312 (.044)	.08		.553 (.022)
Manuf., nondurables	.10		.301 (.044)	.09		.493 (.022)

Table 1 (continued)

Variables	Men: Means	Men: Regression Coefficients: Simple	Men: Regression Coefficients: Multiple	Women: Means	Women: Regression Coefficients: Simple	Women: Regression Coefficients: Multiple
Trans., pub. utils., RR	.01		.394 (.049)	.001		.669 (.078)
Other trans.	.04		.348 (.045)	.02		.575 (.029)
Other utilities	.04		.332 (.045)	.02		.643 (.027)
Wholesale trade	.05		.253 (.045)	.02		.486 (.026)
Retail trade	.14		.114 (.043)	.20		.302 (.018)
Fin., ins., real estate	.04		.269 (.045)	.07		.440 (.022)
Misc. serv.--business & repair	.04		.155 (.045)	.02		.390 (.026)
Pers. services ex. private HH	.01		-.008 (.048)	.03		.327 (.024)
Entertainment & recreation	.01		.083 (.049)	.01		.347 (.034)
Medical except hospitals	.01		.372 (.052)	.06		.436 (.022)
Hospitals	.02		.216 (.047)	.08		.501 (.021)
Welfare & religious	.01		-.205 (.049)	.02		.254 (.026)
Education	.06		.125 (.046)	.15		.307 (.022)

Table 1 (continued)

Variables	Men			Women		
	Means	Regression Coefficients		Means	Regression Coefficients	
		Simple	Multiple		Simple	Multiple
Other prof. services	.02		.267 (.047)	.02		.482 (.026)
Forestry & fisheries	.003		.250 (.067)	.001		.433 (.092)
Public admin.	.07		.347 (.046)	.05		.501 (.025)
Fraction part-time	.11		-.066 (.032)	.21		.002 (.027)
General educ. development	3.63		.037 (.008)	3.56		.106 (.009)
Specific voc. preparation	5.37		.033 (.004)	4.73		.015 (.005)
Strength	2.48		-.051 (.008)	1.98		.013 (.006)
Physical demands	1.90		-.020 (.005)	1.56		.011 (.005)
Environment	.74		.005 (.005)	.26		-.037 (.007)
R^2		.021	.452		.019	.390
N		24,056	24,056		19,412	19,412

Table 2. Decomposition of the Influence of Control Variables on the Estimation of γ

	Men	Women
Total Effect = $\hat{\gamma}-\tilde{\gamma}$	.175	.153
<u>Decomposition</u>		
Schooling and experience	-.024	.024
Region	-.002	-.001
SMSA variables	-.015	-.001
Race	.003	.001
Part-time status	.019	.009
Marital status	.031	.002
Children	.003	.001
Union status	.029	.022
Government employment	.015	.001
Industry	.152	.083
Occupational characteristics	-.034	.012

Table 3. Estimates of γ (with Standard Errors) in the Public and Private Sectors

	Men		Women	
	Public	Private	Public	Private
Simple regression	-.053	-.456	-.271	-.234
	(.031)	(.017)	(.027)	(.014)
Multiple regression	-.203	-.151	-.137	-.094
	(.033)	(.017)	(.027)	(.017)
Multiple regression without industry variables	-.281	-	-.198	-
	(.032)		(.025)	
$\bar{W}$	1.823	1.742	1.531	1.291
$\bar{F}$	.27	.20	.71	.71
N	4,219	19,837	4,440	14,972

Appendix

The full regression model is

$$W = \gamma F + Z\beta + \epsilon.$$

In what follows, it will be convenient to eliminate the constant term by interpreting all variables as deviations from sample means. The simple regression estimator of γ, omitting Z, is then

$$\tilde{\gamma} = (F'F)^{-1}F'W.$$

We wish to prove that the multiple regression estimator of γ, with Z included, is

$$\text{(A1)} \quad \hat{\gamma} = \tilde{\gamma} - b'\hat{\beta}$$

where

$$b = Z'F(F'F)^{-1}$$

is the vector of coefficients from auxiliary simple regressions of each Z variable on F. Equation (A1) is equivalent to equation (4) in the text.

The multiple regression estimator of the full parameter vector is

$$\begin{pmatrix}\hat{\gamma}\\ \hat{\beta}\end{pmatrix} = \left[\begin{pmatrix}F'\\ Z'\end{pmatrix} (F \ Z)\right]^{-1} \begin{pmatrix}F'\\ Z'\end{pmatrix} W$$

$$= \begin{pmatrix}F'F & F'Z\\ Z'F & Z'Z\end{pmatrix}^{-1} \begin{pmatrix}F'W\\ Z'W\end{pmatrix} .$$

Applying Theil's (1971, p. 18) equation (2.15) for inverses of partitioned matrices shows that the one in the above expression can be written as

$$\begin{pmatrix} A & B \\ B' & C \end{pmatrix}$$

where

$$A = (F'F)^{-1}+(F'F)^{-1}F'Z[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'F(F'F)^{-1},$$

$$B = -(F'F)^{-1}F'Z[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1},$$

and

$$C = [Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}.$$

Then

$$\begin{pmatrix} \hat{\gamma} \\ \hat{\beta} \end{pmatrix} = \begin{pmatrix} AF'W + BZ'W \\ B'F'W + CZ'W \end{pmatrix}.$$

Substitution of the full expressions for B and C into the expression for $\hat{\beta}$ yields

$$\hat{\beta} = -[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'F(F'F)^{-1}F'W$$
$$+ [Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'W.$$

Similar substitutions for A and B in the expression for $\hat{\gamma}$ yield

$$\hat{\gamma} = (F'F)^{-1}F'W + (F'F)^{-1}F'Z[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'F(F'F)^{-1}F'W$$
$$- (F'F)^{-1}F'Z[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'W$$

$$= \tilde{\gamma} + b'\{[Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'F(F'F)^{-1}F'W$$
$$- [Z'Z-Z'F(F'F)^{-1}F'Z]^{-1}Z'W\}$$

$$= \tilde{\gamma} - b'\hat{\beta}.$$

This completes the proof.

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