Male-Female Supply to State Government Jobs and Comparable Worth

Author(s): Peter F. Orazem and J. Peter Mattila
Source: *Journal of Labor Economics*, Vol. 16, No. 1 (January 1998), pp. 95-121
Published by: The University of Chicago Press on behalf of the Society of Labor Economists and the NORC at the University of Chicago
Stable URL: http://www.jstor.org/stable/10.1086/209883
Accessed: 27-10-2016 17:25 UTC

Male-Female Supply to State Government Jobs and Comparable Worth

Peter F. Orazem, *Iowa State University*

J. Peter Mattila, *Iowa State University*

The proportion of women in state government jobs and applicant pools is well explained by a model emphasizing supply-side factors. Relative to men, women's supply is least sensitive to wages in predominantly male jobs and most sensitive to wages in predominantly female jobs. These results suggest that comparable worth policies that shift relative pay toward traditionally female jobs and away from traditionally male jobs will increase the proportion of females in male-dominated, female-dominated, and total state government jobs. The implication is that supply side responses need not prevent comparable worth pay adjustments from raising total female compensation.

I. Introduction

After one-half century of nearly unchanging occupational segregation by sex, women began moving into nontraditional jobs in the 1970s. However, despite evidence of declining differences in the occupational distri-

We thank Kelly Cordaro, Dorothy Sally, and Alex Turk for excellent research assistance and Sue Streeter and Donna Otto for help in preparing the manuscript. Harvey Lapan provided helpful suggestions. We also thank numerous employees in the state of Iowa and personnel professionals in state governments around the United States who assisted us in compiling the data. Partial support from National Science Foundation grant no. 8909479 is gratefully acknowledged.

[*Journal of Labor Economics,* 1998, vol. 16, no. 1]

0734-306X/98/1601-0006$02.50

butions of men and women, the U.S. labor market is still highly segregated. Looking at occupations within establishments reveals even higher degrees of job segregation (Bielby and Barron 1984). The persistence of job segregation has come to dominate the discussion of policies to redress disparate labor market outcomes concerning employed men and women. Wage differences between men and women are smaller within narrowly defined occupations than they are across occupations. Both men and women in predominantly female occupations have lower average pay than men and women in predominantly male occupations. As a result, much of the wage differential between men and women is tied to the different employers and occupational labor markets that men and women inhabit and not to disparate treatment within given employers. Blau (1977) finds that men and women are employed in different firms even within an occupation. Groshen (1991) reported that one-half to two-thirds of the pay gap between the sexes in manufacturing and service jobs could be attributed to differences in occupational status. This has led to allegations of market pay discrimination against predominantly female jobs in addition to allegations of access barriers.

Based on this presumption, some have proposed government intervention in the setting of relative pay between male and female jobs. These comparable worth policies have been relegated to the public sector in the United States, but they have been extended (in various guises) to the private sector in other countries, most notably in Australia and the Province of Ontario in Canada. Underlying comparable worth is the notion that as women's labor supply has increased, they have been crowded into the relatively few sectors in which they can compete for jobs. Thus, occupational barriers imposed by discriminating firms have resulted in both occupational segregation and artificially lower pay for female jobs.[1] The presumption of comparable worth is that, at least in the short run, occupational supply behavior of women and men do not respond to changes in relative wages. Without any supply response, these pay policies can raise female compensation in traditionally female jobs and, thus, raise the overall pay of women relative to men.

Other studies have emphasized supply-side explanations for occupational segregation. Some studies have pointed to differences in expected career length or to intermittent employment spells as explaining differences in incentives for men and women to select occupations. If there are different costs associated with time spent out of the labor force (loss of wage growth or lost skills), women may opt for occupations with low

[1] England (1982), Beller (1982), and Bergman (1989) present arguments and evidence supporting the view that occupational segregation is a result of firm discrimination.

exit costs.[2] Other supply-side explanations of the difference in occupational status by sex have centered on presumed difference in tastes for nonpecuniary job attributes between the sexes (Filer 1983).

If the relative number of men and women in an occupation does vary with supply-side behavior, comparable worth policies aimed at raising compensation to women in traditionally female jobs may have some unexpected consequences. Although there are not comparable estimates for females, there is considerable evidence that male occupational supply elasticities are large.[3] If men are more sensitive than women to occupational wage changes, then men will have an incentive to enter traditionally female jobs if pay for those jobs is increased. Similarly, if pay for traditionally male jobs falls due to these policies, men will move out of these jobs in greater proportions, increasing the proportion of women in traditionally male jobs. Even if women have perfectly inelastic occupational supply curves, elastic male occupational supply curves would suggest that the sex composition of jobs is not invariant to relative wages. Yet studies of the effect of comparable worth on male and female pay implicitly have assumed that the current distribution of men and women across occupations is invariant to changes in relative wages.[4]

This study estimates relative male and female wage elasticities to state government jobs using two data sets. One sample includes information on employees in states that have implemented comparable worth pay plans, states using other factor point plans to set pay, and states that have pay systems that are not set by factor points.[5] In addition, we study data on applicants to Iowa state government jobs. We estimate how the relative

[2] Polachek (1981) and McDowell (1982) present evidence that there are different costs of interrupted work careers between traditionally male and traditionally female careers. Sandell and Shapiro (1980) found that anticipated labor force attachment affected earnings growth for young women.

[3] Freeman (1987*a*) and Orazem and Mattila (1991) provide estimates of male occupational supply elasticities that are highly elastic. Comparable supply elasticities for women have not been published to date.

[4] Sorenson (1990) reviews a large number of studies that measure how the proportion female in an occupation alters occupational pay. The presumption is that raising pay to predominantly female occupations would then reduce pay differences between men and women overall. The potential for men to enter traditionally female jobs is not factored into these projections.

[5] A factor point system evaluates the relative value of each job within a pay system by assigning points to each job. These points are based on the levels of skill, effort, responsibility and working conditions associated with each job. The summed points generate a hierarchy of jobs that are then mapped into relative pay, usually with some adjustments to take into account market pay. Comparable worth plans differ from traditional factor point plans in that market wages are explicitly excluded from the setting of relative pay.

number of female incumbents or applicants responds to changes in market and state government occupational pay.

The concentration on state government labor markets offers many advantages. The most obvious is that state governments have been the focus of most of the effort to implement comparable worth policies. A second advantage is that relative pay across occupations within each state government tends to be fixed across time by the pay plan and does not (necessarily) respond to changing market supply factors (Kim 1989). This implies that relative government wages tend to be exogenous within each state.

In the next section, we provide a theory of occupational supply of men and women to the public sector. This motivates the empirical strategy outlined in Section III. Section IV summarizes the nature of our two data sets and the estimates of our model, along with implications and discussion of alternative demand-side interpretations. Section V briefly states our major conclusions.

II. Theory

We decompose the choice of whether or not to select a public-sector job into two parts. First, an individual selects an occupation. All employers may participate in this occupational market, but each employer is too small to affect market wages for the occupation. Once an individual selects an occupation in stage 1, the individual opts for either a public- or private-sector job in stage 2. Our empirical work will concentrate on the stage 2 decision, but we will first describe the two stages separately.

A. Stage 1: Occupational Choice

The human capital investment theory of occupational choice presumes that an individual selects an occupation so as to maximize expected utility. Utility depends on the earnings in the occupation and the nonpecuniary returns associated with job-specific amenities or disamenities. Because individuals differ in talents and in tastes, not all individuals will rank occupations similarly in terms of expected utility. As a result, an individual j will select occupation i among N alternatives so as to maximize expected utility U_{ij}^e. The form for expected utility is assumed to be

$$U_{ij}^e = U_{ij}(W_{ij}, Z_{ij}), \tag{1}$$

where W_{ij} is individual j's wage in occupation i and Z_{ij} is a vector of occupational amenities and disamenities associated with occupation i.

Due to differences in abilities and locations, the W_{ij} and Z_{ij} will vary across individuals. Individual valuation of those wages and amenities will also differ. Consequently, individuals will rank occupations differently.

The variation across individuals in expected utility across N occupations will generate a distribution of I individuals across N occupations such that

$$I = \sum_{i=1}^{N} L_i, \tag{2}$$

where L_i is the number of agents supplying labor to occupation i.

Differences in U_{ij}^e, W_{ij}, and Z_{ij} across the sexes will cause the proportion of men and women to differ across occupations. Males and females may differ in expected earnings or amenities in an occupation, and they may have different relative valuations of these occupational attributes.

Let M_i be the number of men and F_i be the number of women in occupation i. Women's share of occupation i, S_i^F, is

$$S_i^F = F_i/L_i = S(W_i, Z_i), \tag{3}$$

where total labor supply to occupation i is $L_i = M_i + F_i$. The factors that differ across males and females within an occupation would also be expected to differ across spatially distinct labor markets. As a result, S_i^F is likely to differ across markets.

B. Stage 2: Public- versus Private-Sector Choice

By their nature as multiproduct producers, state governments employ workers from many different occupational labor markets. Governments must compete for workers against other firms in the occupational labor market. Therefore, compensation for public employees must be such that expected utility from work in the government must equal or exceed that offered by other firms in the same occupational market for enough workers to fill public jobs.

If government jobs were identical in all respects to private-sector jobs, then the occupational distribution of public-sector men and women would be identical to that in the private sector. However, considerable evidence supports the presumption that public- and private-sector jobs are not identical in pecuniary and nonpecuniary attributes. On average, public-sector employees receive higher pay than similarly skilled private-sector workers. The pay gap is largest at the federal level, with more moderate pay advantages for state and local workers (Smith 1977; Freeman 1987*b*). These favorable wage gaps are largest for female public-sector workers (Krueger 1988). In addition, employee benefits and job security appear to be more generous in the public sector (Quinn 1982; Freeman 1987*b*). While these nonwage attributes will be very similar for men and women within a given state government pay system, men and women may differ in their tastes for these nonpecuniary job attributes

(Blank 1985). As a result of these gender differences in government pay versus market pay and potential differences in tastes for government job attributes, the relative supply of female and male applicants to a government job may deviate from relative supply to the occupation as a whole.

Given the proportion of women in the occupation, S_i^F, the relative incentive to select public-sector employment depends on public-sector wages, w_i, and public-sector amenities, z_i. The proportion of women in government occupation i can be characterized by

$$s_i^f = \frac{f_i}{m_i + f_i} = s(S_i^F, w_i, z_i). \tag{4}$$

The overall proportion female S_i^f should be positively related to s_i^f since one would expect a larger proportion in the applicant pool and because of gender differences in job-specific skills. If the public sector is small relative to the total occupational labor market, S_i^F can be treated as exogenous to s_i^f. However, if the proportion of women in the occupation is altered by public-sector wages and job attributes, then the exogeneity assumption is suspect. In that case, equation (3) can be used to predict S_i^F so that

$$s_i^f = s(\hat{S}_i^F, w_i, z_i), \tag{5}$$

or in reduced form,

$$s_i^f = s(W_i, Z_i, w_i, z_i). \tag{6}$$

Our main interest is in establishing how the proportion of females in a public-sector job is affected by public-sector wages. The effect of wages on the proportion of females in public employment is not obvious because men and women may differ in their relative sensitivity to government wages. In equations (4)–(6), the derivative of s_i^f with respect to w_i is

$$\begin{aligned} ds_i^f/dw_i = {} & [(m_i + f_i)(df_i/dw_i) - f_i((dm_i/dw_i) \\ & + (df_i/dw_i))]/(m_i + f_i)^2, \end{aligned}$$

where m_i and f_i are the number of men and women in government occupation i. Combining terms, we obtain

$$ds_i^f/dw_i = [m_i(df_i/dw_i) - f_i(dm_i/dw_i)]/(m_i + f_i)^2. \tag{7}$$

The relation in (7) can be converted to elasticity form by multiplying and dividing the right-hand side by m_i, f_i, and w_i. The equation is then written as

$$ds_i^f/dw_i = (\varepsilon_i^f - \varepsilon_i^m)(s_i^f s_i^m)/w_i, \tag{8}$$

where ε_i^j is the wage elasticity of supply to government occupation i for sex j, and s_i^j is the proportion of occupation i incumbents who are of sex j. Because an increase in w_i must increase expected utility from choosing job i, ε_i^j must be positive for $j = m, f$. Provided that there is at least one male and at least one female incumbent, $(s_i^f s_i^m)/w_i$ is positive. Therefore, the effect of an increase in government wages on female employment in an occupation will depend on the relative size of the female and male public-sector labor supply elasticities in the occupation. If, for example, women are more sensitive to increases in public-sector wages, then $\varepsilon_i^f > \varepsilon_i^m$ and $ds_i^f/dw_i > 0$. But if the male and female supply elasticities are identical, $(\varepsilon_i^f = \varepsilon_i^m)$, then (8) implies that the coefficient on the public-sector wage will be zero.

The theory suggests that one can determine the relative size of male and female wage elasticities to government employment by estimating a regression equation approximating equations (4)–(6). The most obvious problem with this strategy is that the number of incumbents in a job is a function of both labor demand and labor supply. Therefore, the wage elasticity will reflect both demand and supply effects. However, the concentration on the proportion female in the job rather than the total number of incumbents as the dependent variable makes it possible to sidestep some of these problems. Let $l_i = f_i + m_i$ be the total number of public-sector positions in job i, where m_i is the number of male incumbents and f_i is the number of female incumbents. Government labor demand decisions will set the number of positions (l_i). If governments do not discriminate by sex in hiring, the proportion of female incumbents to the total will reflect the proportion of female incumbents in the qualified application pool. In this way, $s_i^f = f_i/l_i$ will reflect the relative incentives of males and females to supply labor to government. In contrast, if governments systematically discriminate against women, the proportion of women in government jobs will reflect demand-side tastes for discrimination as well as relative supply-side reactions to wages. We will examine the demand- versus supply-side explanations in the empirical section.

III. Empirical Strategy

Equations (4)–(6) suggest that one can derive an estimate of the relative size of male and female occupational supply elasticities by estimating regressions of the form

$$s_i^f = \alpha_0 + \alpha_1 S_i^F + \alpha_2 w_i + \alpha_3 z_i + e_i^A, \tag{9a}$$

$$s_i^f = \beta_0 + \beta_1 \hat{S}_i^F + \beta_2 w_i + \beta_3 z_i + e_i^B, \tag{9b}$$

and

$$s_i^f = \gamma_0 + \gamma_{1w} W_i + \gamma_{1z} Z_i + \gamma_2 w_i + \gamma_3 z_i + e_i^c, \tag{9c}$$

where e_i is an error term and the other variables are defined above.

In (9a), the proportion of female incumbents in the market as a whole is assumed to be exogenous to the proportion of female incumbents in state government. If this assumption is invalid, the instrumental $\hat{S}_i^F$ or its reduced form is preferred. As will be shown, the qualitative results for the coefficient on w_i are not sensitive to the specification choice.

Equations (9a)–(9c) presume that the starting pay for each job is predetermined. This assumption is quite reasonable, especially when pay is measured in relative terms. Government pay structures tend to be rigid: pay increases are generally implemented across the board. A study of the California merit pay system by Kim (1989) found that relative pay in the 1980s could be predicted with near certainty by relative pay in the 1930s. This institutional rigidity implies that starting wages in the merit system are not altered in response to changes in the sex composition of incumbents in the job. It is, of course, possible that these relative wages were sex-biased when set years ago, but for our purposes, it is only important that they be viewed as predetermined data by workers.

An exception to this argument lies in the implementation of comparable worth pay plans. Since comparable worth states base pay at least partly on proportional female incumbency and because of their recent implementation, predetermination of state relative pay using 1987–88 data may be questionable. Fortunately, we also have information from states that have not implemented comparable worth systems. We test the government pay exogeneity assumption by examining the sensitivity of the parameter estimates when only noncomparable worth states are included. Their pay systems have been in place longer and have no explicit goal to link pay to the proportion of female incumbents.

A second way to remove possible simultaneity between s_i^f and w_i is to use the proportion of female applicants for job vacancies rather than the proportion of female employees as the dependent variable. Because pay is set before jobs are advertised, applicants must be responding to public-sector wages. As we show below, results using a sample of applicants were similar (although not identical) to those using the sample of incumbents.

IV. Data and Results

A. Multistate Analysis

In this section we investigate data on incumbents gathered from a survey of state government personnel departments. After obtaining a complete listing of job titles from a subset of states, we selected 78 job titles that were common to most states, that matched published *1980 Census of Population* occupation titles (U.S. Bureau of the Census 1983, table 222), and that spanned the range of traditionally male, traditionally female, and mixed jobs such as carpenter, clerk, and registered nurse. The list of jobs by gender composition is included in the appendix. Phone interviews with state personnel professionals indicated that these jobs were also broadly representative of the range of jobs performed in state government. State personnel managers were asked to match the job title to the closest position in their state pay system. Respondents were asked to call for clarifications if they were uncertain about the job title, and titles that were not easily recognized were dropped. For each job, the personnel professional was asked to report the starting pay, the proportion female, and whether the job was covered by a union contract. In addition, information was obtained on whether the state's pay system was a comparable worth pay system, a factor point pay system, or some other pay system.

Sixteen states declined to participate or did not have consistent statewide pay systems. Another 14 had information on pay and union status but did not have information on female incumbents by job classification. The remaining 20 states supplied the necessary information. These states are listed in table 1 along with information on their pay system and public-sector unionization. Pooling across the 20 states, we had complete information on 1,393 state jobs, an average of almost 70 job titles per state pay system. The information provided was from state pay systems in 1987 or 1988.

Seven of the 20 state systems are not covered by collective-bargaining agreements, but workers in states with comparable worth systems are more likely to be covered by a collective bargaining agreement. The information in table 1 also shows that in 16 of the 20 states, the proportion of women incumbents in state government is higher than their relative proportion in the state labor market as a whole. Several states, including Connecticut, Delaware, Nevada, North Carolina, and Virginia, have much higher levels of female state government incumbency than in the same occupations in the state as a whole.[6]

[6] Our state occupation data are taken from 1980 census figures (U.S. Bureau of the Census 1983). Despite the 8-year gap between the census data and our survey data, our results are quite consistent with national data for 1988. Averaging across all jobs in our multistate sample, women represented an average of 50.5% of state incumbents as opposed to 45.7% of employees in the related occupation in the state. These numbers are comparable to 1988 annual averages as reported in the U.S. Bureau of Labor Statistics' (1989) *Employment and Earnings.* Women represented 49.1% of all public employees and 45.7% of employees in private establishments.

Table 1
Characteristics of State Pay Systems: Multistate Sample of Incumbents

	Job Titles	Union	s^f	S^f
Comparable worth states:				
Connecticut	67	.94	.54	.45
Iowa	77	.70	.49	.48
Michigan	57	.90	.54	.47
Minnesota	75	.89	.50	.47
Oregon	72	.96	.51	.49
Washington	78	.00	.40	.47
Wisconsin	75	.88	.45	.46
Other factor point states:				
Delaware	64	.53	.55	.46
Maine	67	.97	.47	.42
New Hampshire	76	.94	.55	.47
Oklahoma	78	.00	.45	.48
Vermont	69	1.00	.50	.53
Wyoming	67	.00	.56	.47
Nonfactor-point states:				
Alaska	63	.98	.46	.44
Colorado	69	.00	.51	.44
Montana	65	1.00	.55	.48
Nebraska	72	1.00	.46	.45
Nevada	61	.00	.49	.38
North Carolina	71	.00	.56	.43
Virginia	74	.00	.58	.46

NOTE.—Union, s^f, and S^f are unweighted averages across job titles within each state. The Union column reports the proportion of job titles covered by collective bargaining in the state pay system; figures in the Union and s^f columns are based on information provided by a survey of state government personnel specialists; figures in the S^f column were computed from 1980 U.S. census state-level data on all men and women in the detailed occupations that most closely resemble the state government job titles. Some of the comparable worth states such as Michigan have made some comparable worth pay adjustments without instituting a complete study or adjustment. Washington implemented a comparable worth plan only after a legal suit was dropped as a result of union negotiations.

The measures of proportion female and union coverage are self-explanatory. The measure of state occupation pay (w) requires further comment. Differences in cost-of-living, benefits, job conditions, local amenities, or other factors may alter the relative attractiveness of state government jobs across states. To control for variation in these unmeasured compensating differentials, we normalize pay in state job i relative to an "average" pay rate computed for each state. There were eight occupations that were common to all states for which we had pay information.[7] An average base was computed across these eight occupations by state using

$$w_{ij} = ((1/8) \sum_{i=1}^{8} w_{ij})^{-1} w_{ij}, \tag{10}$$

[7] These eight occupations included computer operator, cook, laundry worker, librarian, library assistant, licensed practical nurse, occupational therapist, and social worker.

This measure, together with a set of state dummy variables, removes the biases associated with interstate compensating differentials.

The use of publicly known, institutional pay plans with clear entry wage rates implies that males and females will receive the same wage rate. Nondiscrimination policies and public scrutiny assured this within the public sector in the late 1980s. As Blau (1977) and Groshen (1991) imply, private-sector wages may differ by gender within an occupation because of gender segregation and wage differentials between firms. In an attempt to control for this, we compute the ratio of female-to-male earnings within the state (W^F/W^M), which becomes an additional control variable.

Table 2 contains definitions and sample statistics of the variables used in the analysis of the 20-state incumbent data set. The sample statistics are reported for all jobs and for predominantly male and predominantly female jobs. Male jobs were defined (following much of the literature) as jobs in which, when averaged across all 20 state governments, more than 75% of all incumbents were male. These represent about 30% of all jobs in the sample. Similarly, female jobs were defined as jobs in which more than 75% of the incumbents are female. Twenty-six percent of jobs in the sample were predominantly female, leaving 44% of the jobs implicitly defined as mixed. The average male-dominated job was 90% male, and the average female-dominated job was 88% female.

The measure of relative female status in each occupation in the state as a whole comes from the *1980 Census of Population* data on year-round full-time workers in detailed occupations. The measure is based on the number of men and women in the state whose occupation most closely matches the job title in the state government. The 1980 census was used to reduce the likelihood that S_{ij} and s_{ij} were simultaneously determined.

A comparison of the proportion female in state government jobs and the proportion female in the state reveals considerable consistency. Just over half of the state government employees in these occupations are female, whereas 46% of the workers in these occupations in the state as a whole are women. In both traditionally male and traditionally female occupations, the proportion female in state government jobs is about 2% higher than the proportion female in the state as a whole.

As discussed above, S_i^F, the proportion of women in the occupation as a whole may be endogenous if the public sector hires a large proportion of the relevant occupational supply. Following equation (3), a vector of instruments is used for S_i^F. One of the variables is (W^F/W^M), the annual income of full-time women relative to full-time men in the occupation as a whole. The other instruments are measures of overall occupational attributes, Z_i, including average occupational pay, W, required previous job experience, required ability in reading, mathematics and logic, required strength, and estimates of hourly benefits in the occupation. The required experience, ability, and strength measures come from the *Dic-*

Table 2
Sample Statistics and Definitions of Variables: Multistate Sample of Incumbents

	All Jobs	Male Jobs	Female Jobs
Dependent variable:			
s^f: proportion female in government job	.505	.101	.888
	(.37)	(.16)	(.16)
Government job variables:			
w: wage in government job (normalized)*	1.103	1.228	.930
	(.36)	(.48)	(.20)
Male Job: dummy variable indicating >75% male incumbents	.296	. . .	. . .
	(.47)		
Female Job: dummy variable indicating >75% female incumbents	.261	. . .	. . .
	(.44)		
Union: proportion of workers in government job covered by collective bargaining	.577	.594	.567
	(.49)	(.49)	(.49)
Market variables:			
S^F: proportion female in reference occupation	.457	.080	.865
	(.34)	(.09)	(.143)
W: earnings in reference occupation (normalized)*	1.383	1.723	1.006
	(.64)	(.91)	(.218)
W^F/W^M: relative female to male earnings	.722	.679	.772
	(.18)	(.21)	(.19)
Required Experience (months)	30.32	34.79	18.22
	(27.0)	(28.5)	(19.3)
Required Reading	4.09	4.01	3.72
	(1.06)	(1.20)	(.78)
Required Mathematics	3.14	3.05	2.71
	(1.25)	(1.33)	(.72)
Required Logic	3.68	3.25	3.59
	(1.25)	(1.39)	(.91)
Required Strength	2.15	2.80	1.79
	(.89)	(.89)	(.76)
Benefits	1.45	1.43	1.22
	(.50)	(.53)	(.42)

NOTE.—Unweighted averages across all states and jobs with standard deviations in parentheses. The dependent variable and the government job variables are from the Survey of state government personnel specialists. The market variables S^f, W, and W^F/W^M are from the 1980 U.S. census state-level data on detailed occupations. The "Required" variables are from the *Dictionary of Occupational Titles* (U.S. Department of Labor 1977), and the "Benefits" variable is from the U.S. Bureau of Labor Statistics (various issues).
* This is normalized relative to the "average" state wage defined in eq. (10).

tionary of Occupational Titles (U.S. Department of Labor 1977), while the benefits data are compiled from "Employer Costs for Employee Compensation—March" (U.S. Bureau of Labor Statistics, various issues) for broad occupational groups. Predominantly male jobs disproportionately require experience and strength. On average, predominantly female jobs require less reading and mathematics skills but more logic skills than predominantly male jobs. Female jobs have lower benefits than the average across all jobs.

The generalized least squares estimates of equations (9a)–(9c), explaining the proportion of female state government incumbents by detailed occupation in the 20 states, are reported in tables 3 and 4.[8] All regressions include the unionization variable and state dummy variables as elements of the vector z_i.

We interact government starting pay (w) with dummy variables for male- and female-dominated jobs. This allows inferences to be drawn on how wage policies aimed at raising relative pay for female-dominated jobs will affect the sex composition of state jobs. The coefficient on (uninteracted) w_i represents the response of the proportion of women incumbents to changes in wages in mixed gender jobs. These coefficients are always positive, indicating that women are more sensitive to state government pay in mixed-gender jobs than are men. All of the coefficients on government wages interacted with the female-dominated job dummy variable are positive and significant. The implication is that women's labor supply to predominantly female jobs is relatively more elastic than their supply to mixed jobs. The total effect (determined by adding the coefficients on w and $w*$Female Job) is positive as well. The reverse is true for predominantly male state jobs. The sum of coefficients on w and $w*$Male Job are negative, indicating that men have higher elasticities of supply to predominantly male jobs than do women.

These results suggest that a policy to raise relative pay for predominantly female public jobs would raise the proportion of women in female dominated and mixed jobs. If, as is likely, such increases for predominantly female jobs come at the expense of raises in traditionally male jobs, then relative wages in the latter jobs will drop. Because women have lower supply elasticities in male-dominated jobs, the proportion of women in traditionally male jobs will also increase as relatively more males exit these jobs. Therefore, these results suggest that a comparable worth pay policy would tend to raise the proportion of women employed in both female-dominated and male-dominated state government jobs.

Because hiring policies in government are typically based on formal tests and civil service procedures across all occupations, we minimize the likelihood that our results are due to discriminatory hiring policies. Equally important, government hiring in the 1980s was subject to much public scrutiny and political pressure to avoid discrimination. If the public sector has been less discriminatory than private-sector firms in recent

[8] Because of the potential massing of observations on the dependent variable at zero and at one, we also estimated these equations (and those in table 7) using a Tobit specification with upper and lower bounds. None of the findings changed. Examination of the error terms indicated that the least squares specification was less subject to heteroskedastic errors than was the Tobit specification. Standard errors were corrected for heteroskedasticity, using White's (1980) method.

Table 3
Regression Analysis of Female State Government Occupational Shares in 20 States

	Equation		
	(9a)	(9b)	(9c)
w	.093**	.143**	.111**
	(3.21)	(4.50)	(2.36)
w*Female Job	.167**	.309**	.355**
	(6.36)	(17.2)	(16.6)
w*Male Job	−.187**	−.292**	−.291**
	(9.23)	(14.2)	(14.6)
Union	.011	.066	.087*
	(.27)	(1.48)	(1.90)
Market variables:			
S^F	.502**		
	(12.2)		
$\hat{S}^F$		.278**	
		(9.21)	
W			.034
			(1.49)
W^F/W^M			.015
			(.37)
Required Experience			−.001
			(1.28)
Required Reading			−.074**
			(2.86)
Required Mathematics			.035**
			(2.90)
Required Logic			.027
			(1.34)
Required Strength			−.070**
			(6.38)
Benefits			−.031
			(1.38)
R^2	.65	.61	.61
N	1,392	1,392	1,392

NOTE.—All regressions include 19 dummy variables for each state and an intercept. t-statistics corrected for heteroskedasticity are reported in parentheses. The dependent variable is the natural logarithm of S^F.
* Significant at the 10% level.
** Significant at the 5% level.

years, it is less likely that our results are attributable to employer hiring discrimination.

However, if some state employers have tastes for discrimination for or against women, the estimates in table 3 will reflect demand as well as supply-side decisions. Given voluntary enactment, comparable worth states would be least likely to have tastes for discrimination against women and might even discriminate against men. This is especially true since our data set was collected after comparable worth had been implemented in these states. Conversely, the noncomparable worth states would be more likely to have discriminatory tastes against women. To

Table 4
Regression Analysis of Female State Government Occupational Shares by State Pay System

	Equations								
	Comparable Worth States			Other Factor Point States			Not Factor Point States		
	(9a)	(9b)	(9c)	(9a)	(9b)	(9c)	(9a)	(9b)	(9c)
w	.054	.138**	.109	.116**	.170**	.020	.108**	.134**	.206**
	(1.24)	(3.20)	(1.54)	(2.25)	(2.61)	(.23)	(3.29)	(3.17)	(3.01)
w * Female Job	.126**	.339**	.380**	.177**	.275**	.336**	.183**	.308**	.345**
	(3.98)	(7.97)	(13.8)	(4.16)	(12.7)	(9.97)	(5.18)	(11.6)	(13.0)
w * Male Job	−.125**	−.275**	−.284**	−.202**	−.282**	−.256**	−.222**	−.319**	−.329**
	(5.58)	(9.60)	(10.5)	(5.72)	(12.8)	(8.02)	(8.66)	(13.6)	(12.8)
Union	.007	.072	.085**	−.036	.040*	.090	.386**	.513**	.503**
	(.23)	(.52)	(2.17)	(.49)	(1.82)	(1.25)	(7.57)	(11.3)	(8.63)
Market variables:									
S^F	.592**			.418**			.526**		
	(12.5)			(6.56)			(8.62)		
$\hat{S}^F$		.223**			.322**			.294**	
		(6.95)			(5.28)			(6.96)	
W			.026			.070			−.013
			(.84)			(1.66)			(.44)
W^F/W^M			.009			.007			.009
			(.10)			(.12)			(.19)
Required Experience			−.0006			−.0015*			.0000
			(.95)			(1.76)			(.04)
Required Reading			−.020			−.095**			−.103**
			(.62)			(2.35)			(2.73)
Required Mathematics			.007			.064**			.033*
			(.45)			(3.24)			(1.83)
Required Strength			−.046**			−.098**			−.072**
			(3.32)			(6.07)			(4.84)
Benefits			−.011			−.006			−.077**
			(.41)			(.18)			(2.56)
R^2	.73	.64	.64	.57	.55	.58	.67	.63	.63
N	501	501	501	416	416	416	475	475	475

NOTE.—All regressions include 19 dummy variables for each state and an intercept. *t*-statistics corrected for heteroskedasticity are reported in parentheses.
* Significant at the 10% level.
** Significant at the 5% level.

Table 5
Differences in Supply Elasticities between Men and Women and Simulated Response to Comparable Worth Wage Adjustments

A. Elasticities

	Mixed Jobs	Female Jobs	Male Jobs
1. All states	.63	4.24	−2.01
2. Comparable worth states	.61	3.97	−1.85
3. Other factor point states	.74	3.73	−1.24
4. Not factor point states	.60	5.72	−3.06
5. Iowa sample of applicants to state government jobs	−.69	.30	−2.74

B. Response to 11% Increase in Pay or Predominantly Female Jobs

	% Change in Proportion Female			
	Mixed Jobs	Female Jobs	Male Jobs	Total
6. All states	−1.13	3.06	10.1	3.25
7. Comparable worth states	−1.13	3.34	9.54	3.17
8. Other factor point states	−1.15	3.08	5.13	1.68
9. Not factor point states	−1.16	2.84	2.26	.65
10. Iowa sample of applicants to state government jobs	1.43	.10	9.37	3.42

NOTE.—All estimates are based on the two-stage parameter estimates in tables 3, 4, and 6. In pt. B, the 11% increase in relative pay for predominantly female jobs assumes a 6.4% increase in female jobs and a 4.2% decrease in other jobs, leaving average government wages unchanged.

examine these hypotheses, equations (9a)–(9c) were estimated separately for comparable worth states, factor point states, and non-factor-point states. Non-comparable-worth states were dichotomized into the two latter groups given the possibility that nonfactor point states might have more flexibility to discriminate. The results, reported in table 4, are remarkably consistent across subsamples. If these results are due to demand-side tastes for discrimination, then these tastes are so pervasive as to be similar in the most proactive and traditional states.

To make these results easier to examine, implied differences in public-sector wage elasticities across the sexes are reported in table 5, evaluated at sample means. Because the results are similar across specifications, we report only those using coefficients from estimates of (9b). The computation is performed by transforming equation (8) so that

$$(\varepsilon_i^f - \varepsilon_i^m) = (ds_i^f / dw_i)(w_i / s_i^f s_i^m), \tag{11}$$

where the derivative is the appropriate coefficient from tables 3 and 4.

The difference in elasticities reported in table 5 tell a consistent story across all pay plans. Women are more sensitive to changes in public-sector pay for mixed and female dominated jobs. The gap between wage elasticities is largest in the female-dominated jobs. In contrast, men have much higher wage elasticities in traditionally male jobs. The differences between male and female wage elasticities are larger in states using pay plans that are not based on factor points, perhaps indicating demand-side tastes that reinforce supply-side differences.

If the relative elasticities in table 5 are to be explained by discriminatory selective hiring by sex, then it must be that state governments give preferential treatment to hiring males as occupational entry-level wages rise in predominantly male jobs. At the same time, state governments would have to give preferential treatment to hiring females as the occupational entry-level wage rises in mixed and female dominated jobs. Since none of the positions are supervisory, and all are entry-level positions filled from outside hires, the pattern cannot be explained by preferential promotions within the state job hierarchy. While it is possible that states engage systematically in selective preferential hiring policies that vary by entry-level wage and gender composition of the job, the supply-side story seems more plausible. That is, although employers may prefer males as the wage rises in male dominated jobs, it seems unlikely that they would also prefer females as the wage rises in mixed and female-dominated jobs.

B. Iowa Applications Analysis

To better distinguish supply decisions from demand decisions, it is preferable to use measures of notional supply to state governments, rather than measures of incumbents already employed. Such data were available for 148 jobs in the Iowa state government for the years 1986 ($\frac{1}{2}$ year), 1987–88, 1989–90, and 1990–91. Only entry-level occupations were selected to avoid problems associated with applications to jobs for which prior state experience was a prerequisite.

The Iowa applications data includes only those deemed qualified for the job. Qualifications are determined by education, past experience, licensure, or performance on exams as dictated by the particular requirements of each job. Applications are only taken when openings exist. For some jobs, with many incumbents and high turnover, applications are taken continuously throughout the year. For other jobs, applications may only be taken for a month or two. Therefore, the use of relative numbers of applicants by gender, rather than the absolute number of applicants, immediately controls for the length of time that applications were being accepted. Both males and females have equal opportunity to apply, and the relative number of women in each applicant pool represents the relative supply of women to state government.

The regressors include those used in the multistate analysis of incumbents plus a few other variables that were thought to affect the marginal utility of accepting state employment. The state government wage is measured as the starting wage in the job as taken from state pay plans for the relevant years. The market wage was computed as the average wage from the March Current Population Survey (CPS) tapes (U.S. Bureau of the Census, various years) of the occupation most closely tied to the state occupation for the 3 years preceding the application year. The applications data better matched the more detailed CPS occupational titles as opposed to the more broadly defined published census tables used for the multistate sample. The comparison was made by using the Iowa Department of Personnel's (1986) "Minimum Qualifications Guide" to identify occupations or skills deemed qualifying for entry into the state job. Regional wages were used for jobs in which local markets were presumed to be most relevant. National averages were used for jobs that are primarily highly skilled and for which national recruitment was presumed to occur. The CPS tapes also provided information on the relative number of women and men in the occupation.[9]

These variables were supplemented by several other measures of the attractiveness of the job. The first is the number of openings in the job during the year. This was derived from the December payroll tapes as the number of incumbents in each job with less than 1 year of experience with the state. The potential for advancement was measured by the number of pay grades one could advance from the entry job before becoming a supervisor. This was constructed using the Iowa "Minimum Qualifications Guide," which provided information on jobs for which experience in the entry job would be considered qualifying for internal promotion. That source also provided information on the educational and experience requirements of the job.

The sample statistics for these variables are reported in table 6. Male- and female-dominated jobs were again defined to be those with more than 75% incumbents of that sex. The first finding from table 6 is that applicants do not match the current sex composition of the job. While women make up 7% of the incumbents (as computed from Iowa payroll tapes) in the male jobs, they make up over 20% of the qualified applicants to those jobs. Similarly, men make up 23% of the qualified applicants to female jobs, whereas current incumbents are only 7% male. This would suggest the potential for large changes in the sex composition of these jobs in the future. It also suggests that historical labor supply decisions (as reflected by those already employed) may differ from labor supply decisions made by current job entrants.

[9] The job titles in the Iowa applications data set did not all match census occupations. The CPS tapes have more narrowly defined occupations than do the published detailed census data, but the CPS has much smaller sample sizes. Pooling over 3 years of March Current Population Survey data was necessary to insure large enough samples to estimate average market wages for narrowly defined occupations.

Table 6
Sample Statistics and Definitions of Variables: Iowa State Applications Sample

	All Jobs	Male Jobs	Female Jobs
Dependent variable:			
s^f: proportion female applicants for state jobs	.506	.201	.770
	(.30)	(.20)	(.21)
Government job variables:			
w: starting biweekly wage/100 for government job	7.26	7.89	6.64
	(1.52)	(1.54)	(1.42)
Union: dummy variable indicating if the state job is covered by union contract	.677	.630	.635
	(.47)	(.48)	(.48)
Male Job: dummy variable indicating >75% male incumbents	.302	. . .	. . .
	(.46)		
Female Job: dummy variable indicating >75% female incumbents	.326	. . .	. . .
	(.47)		
Openings: number of vacancies to be filled	33.4	13.3	50.0
	(105.0)	(28.4)	(121.8)
Pay Growth: potential pay grade growth in occupational job ladder	3.51	2.78	3.85
	(3.14)	(3.20)	(3.22)
Required (Iowa) Education (years)	13.9	13.7	13.5
	(2.61)	(2.87)	(1.86)
Required (Iowa) Experience (years)	1.34	1.69	1.45
	(1.47)	(1.51)	(1.48)
Market variables:			
W: 3-year average biweekly wage/100 for reference occupation	8.41	9.30	7.07
	(3.01)	(2.88)	(2.38)
S^F: Proportion female in reference occupation	.531	.266	.776
	(.31)	(.24)	(.21)
W^F/W^M	.697	.685	.703
	(.12)	(.10)	(.13)
Required Experience (months)	35.13	44.81	26.78
	(27.08)	(29.46)	(23.69)
Required Reading	4.41	4.36	4.22
	(.83)	(.80)	(.85)
Required Mathematics	3.47	3.62	3.18
	(1.08)	(1.19)	(.88)
Required Logic	3.91	3.96	3.64
	(1.03)	(1.03)	(.98)
Required Strength	1.82	2.20	1.66
	(.80)	(.85)	(.73)
Benefits	1.56	1.60	1.44
	(.47)	(.46)	(.44)

NOTE.—These are unweighted averages across all detailed jobs within Iowa with standard deviations in parentheses. Variables W and w are deflated by the implicit price deflator for consumer goods, 1987 = 100. The market variables W, S^F, and W^F/W^M are the CPS occupations deemed to provide the best background to meet the requirements of state job i, as defined in the "Minimum Qualifications Guide" (Iowa Department of Personnel 1986). The "Required" variables are from the *Dictionary of Occupational Titles* (U.S. Department of Labor 1977), and the Benefits variable is from the U.S. Bureau of Labor Statistics (various issues).

The sample statistics for the Iowa job attributes are reported in table 6. Starting wages in predominantly female-dominated jobs in Iowa State government are lower than starting wages in predominantly male jobs. Pay in mixed jobs is lower than that in predominantly male jobs but higher than that in predominantly female jobs. Male government jobs require slightly higher levels of education and prior experience than do female jobs. The relative market gap in education and experience requirements between male and female jobs is larger, based on the *Dictionary of Occupational Titles* (U.S. Department of Labor 1977) measures for these jobs. Male government jobs provide less opportunity for pay growth and have fewer openings per year than do female jobs. The former is contrary to the general pattern (Blau 1977, p. 100) and may reflect the greater number of female-dominated medical occupations (nurses, therapists, etc.) found in state government.

The estimates of equations (9a)–(9c) using the sample of qualified applicants in Iowa state government are reported in table 7. The results for predominantly male and female jobs are very consistent with those obtained from use of the multistate data set on incumbents. Female relative wage elasticities are greatest for predominantly female jobs and smallest in predominantly male jobs, as in the sample of incumbents. For mixed jobs, female supply is less elastic or equally elastic compared to men, in contrast to the larger female elasticities obtained for mixed jobs in the sample of incumbents.

The difference in male- and female-wage elasticities are reported in row 5 of table 5. For male-dominated jobs, the estimates are similar to those in the incumbents sample. Males are much more sensitive than women to government wages in traditionally male jobs. For traditionally female jobs, women still have larger wage elasticities relative to men. The gap is much smaller than in the earlier estimates but is statistically significant. Men now have more elastic supply to mixed jobs, but the difference is only marginally significant. In both data sets, the female wage elasticity increases relative to the male elasticity as the analysis moves from male-dominated to mixed to female-dominated jobs.

The larger differences in male and female wage elasticities in the incumbents sample versus the applicants sample is consistent with the presumption that male and female labor supply decisions are becoming more similar over time. Since incumbents would be expected to be from older cohorts on average than are job seekers, one would expect smaller differences in wage elasticities in the applicants sample.

The Iowa applications sample offers a further check on whether the proportion of female incumbents by job reflects largely supply-side decisions. Using the same 148 occupations, we regressed the proportion female among new hires on the proportion female in the applicant pool using data averaged over the available years. The null hypothesis

Table 7
Regression Analysis of Female Iowa State Government Application Shares

	Equation		
	(9a)	(9b)	(9c)
$w/100$	−.017*	−.024*	.002
	(1.67)	(1.66)	(.12)
$w*$Female Job	.018**	.032**	.029**
	(5.46)	(8.70)	(8.07)
$w*$Male Job	−.020**	−.032**	−.030**
	(5.81)	(9.66)	(8.95)
Union	−.016	−.046**	−.070**
	(.76)	(1.99)	(3.07)
Openings	.0104**	.0353**	.0179**
	(2.12)	(4.86)	(3.80)
Required (Iowa) Education	.008	.007	.014*
	(1.75)	(1.10)	(1.91)
Required (Iowa) Experience	−.002	.0005	.002
	(.30)	(.06)	(.26)
Pay Growth	.002	−.006**	−.0003
	(.61)	(1.81)	(.07)
Market Variables:			
S^F	.493**		
	(10.8)		
$\hat{S}^F$		.198**	
		(4.24)	
$W/100$			−.029**
			(4.89)
W^M/W^F			−.094
			(1.00)
Required Experience			−.001**
			(2.30)
Required Reading			−.008
			(.41)
Required Mathematics			.024
			(1.33)
Required Logic			.006
			(.31)
Required Strength			−.067**
			(5.10)
Benefits			−.070**
			(2.95)
Intercept	.267**	.556**	.861**
	(3.29)	(6.16)	(6.99)
R^2	.68	.57	.61
N	415	415	415

NOTE.—t-statistics corrected for heteroskedasticity are reported in parentheses.
* Significant at the 10% level.
** Significant at the 5% level.

that the share of women among applicants and new hires were equal could not be rejected at the .10 significance level. This is consistent with the view that state government did not discriminate in hiring during this period.

A few other results are worth noting. Required strength and experience and benefit levels in the market lower the proportion of women among incumbents and applicants. Educational skills are not statistically significant in the applications data estimates, but the signs are consistent with the incumbents sample estimates. Relative earnings of men and women in the overall occupation also did not have a significant effect on the proportion of women among incumbents or applicants.

C. Simulations

Sorenson (1987) reported that comparable worth wage adjustments increased relative pay for female jobs by 11%. To show how changes in relative pay affect the relative number of women in state jobs, we used the estimates in tables 3, 4, and 7 to predict the proportion female in government jobs if pay for female jobs relative to other jobs increased by 11%, holding average wages in state government fixed.[10] This was accomplished by raising wages for predominantly female jobs by 6.4% and lowering wages for other jobs by 4.2%. While this exercise allows us to isolate the effects of relative pay on supply, holding fixed the ratio of overall government to private-sector pay, it also has some basis in reality. Governmental budget constraints imply that real pay increases for some jobs may require real pay reductions in others, say by holding pay increases below the rate of inflation.[11]

The largest supply response to a comparable worth policy, as shown in the bottom half of table 5, are in predominantly male jobs. The proportion of female incumbents in male jobs is predicted to increase by 10.1%, while the proportion of female applicants rises by 9.4%. The reason is that men leave government more readily when government wages fall because of the higher male wage elasticity in predominantly male jobs. Women's employment share also rises in predominantly female jobs, albeit by a much smaller percentage than the increase in male jobs. For mixed jobs, the simulated responses imply a small reduction in women's share of employment but a small increase in their share of applicants. Overall, this simulated comparable worth policy raised the proportion of women among incumbents by 3.25% and the proportion of women applicants to state government by 3.42%.[12]

[10] The simulated proportion female does not require that only supply-side factors explain the wage elasticities. The simulations will reflect both demand and supply-side factors if discriminatory tastes affect the wages elasticities in table 5.

[11] See Orazem and Mattila (1990) for a discussion of these issues.

[12] These simulated comparable worth effects are partially consistent with the pattern of female incumbency we observe in the multistate sample. The relative supply elasticities estimated from the Iowa sample suggest that lowering relative pay for male and mixed dominated jobs would raise the proportion female in

Table 8
Unweighted Average Proportion Female

	Male Dominated	Mixed Job	Female Dominated
Comparable worth states	.104	.514	.872
Non-factor-point states	.082	.582	.922

V. Conclusions

Using two different data sets, we find strong evidence that the proportion of women in state government jobs is affected by relative occupational pay. In particular, our findings suggest that women have higher supply elasticities in predominantly female jobs and men have larger supply elasticities in male jobs. While our findings might be explained by employment discrimination in favor of women in higher-paid female jobs and in favor of men in higher-paid male jobs, it is more likely that these results are attributable to different public-sector supply elasticities between men and women. The supply-side interpretation is supported by similar findings in a sample of job applicants and evidence supporting the hypothesis of sex-blind hiring from application pools.

The findings suggest that policies that alter relative pay in state government will alter the gender composition of state jobs. In particular, comparable worth policies that raise relative pay in traditionally female jobs while lowering relative pay in traditionally male jobs will raise the proportion of females in male-dominated and female-dominated jobs and will tend to increase the female share of public-sector jobs more generally. The implication is that supply-side responses by themselves need not prevent comparable worth pay adjustments from raising total female compensation. This conclusion must be qualified by noting the potentially offsetting decline in female compensation as women gain an increasing proportion of male-dominated jobs that suffer falling real wages.

male jobs and lower the proportion female in mixed jobs. The prediction is consistent with the multistate, cross-sectional pattern of female incumbency as shown in table 8. However, raising pay in the female-dominated jobs should raise the proportion female in comparable worth states, in contradiction to the observed cross-sectional pattern. It should be emphasized that a better test would be to observe female incumbency longitudinally as relative pay is changed within a state, but that is not possible with the current data.

Appendix

Table A1
Seventy-Eight Jobs in the Multistate Sample of Incumbents

	Number of States Responding	Mean % Female	Job Type
Professional and management related:			
Accountant	21	.615	Mixed
Attorney I	14	.365	Mixed
Audiologist	14	.693	Mixed
Auditor I	20	.456	Mixed
Budget Analyst	19	.341	Mixed
Business Manager I	17	.423	Mixed
Chaplain I	19	.086	M-dom
Chemist I	20	.473	Mixed
Dentist	19	.169	M-dom
Dietitian	19	.884	F-dom
Librarian I	21	.730	Mixed
Microbiologist	21	.582	Mixed
Occupational Therapist I	21	.767	F-dom
Personnel Officer	20	.583	Mixed
Pharmacist I	18	.462	Mixed
Physical Therapist I	19	.607	Mixed
Physician	16	.232	M-dom
Probation and Parole Officer	20	.357	Mixed
Psychologist	21	.241	M-dom
Purchasing Agent	20	.528	Mixed
Registered Nurse I	19	.832	F-dom
Safety Officer I	16	.293	Mixed
Social Worker	21	.738	Mixed
Speech Therapist	19	.734	Mixed
Statistical Analyst I	19	.502	Mixed
Tax Auditor	19	.400	Mixed
Veterinarian	18	.070	M-dom
Vocational Rehabilitation Counselor	19	.543	Mixed
Technical:			
Airplane Pilot	16	.010	M-dom
Dental Hygienist	19	.873	F-dom
Engineering Technician I	20	.153	M-dom
Laboratory Technician	20	.641	Mixed
Licensed Practical Nurse	21	.893	F-dom
Medical Laboratory Technician	16	.689	Mixed
Programmer	21	.432	Mixed
X-Ray Tech	17	.743	Mixed
Operators and laborers:			
Equipment Operator	19	.029	M-dom
Groundskeeper	20	.183	M-dom
Laborer	21	.187	M-dom
Laundry Workers	21	.675	Mixed
Seamstress	16	.868	F-dom
Welder	16	.023	M-dom
Service:			
Baker	16	.358	Mixed
Barber	15	.300	Mixed
Beautician	17	.827	F-dom
Cook	21	.605	Mixed
Correctional Officer I	21	.176	M-dom
Custodial Worker	20	.402	Mixed
Dental Assistant	19	.877	F-dom

Table A1 (*Continued*)

	Number of States Responding	Mean % Female	Job Type
Food Service Worker	20	.600	Mixed
Forest Ranger	17	.116	M-dom
Highway Patrol Officer	18	.081	M-dom
Lab Assistant	20	.710	Mixed
Nursing Aid	18	.764	F-dom
Psychologist Assistant	12	.546	Mixed
Security Guard	20	.188	M-dom
Administrative support and sales:			
Administrative Assistant	20	.863	F-dom
Cashier	15	.844	F-dom
Clerk	20	.848	F-dom
Clerk Typist	21	.949	F-dom
Computer Operator I	20	.503	Mixed
Data Entry Operator I	21	.896	F-dom
Interviewer	14	.641	Mixed
Library Assistant	21	.770	F-dom
Personnel Assistant	21	.931	F-dom
Postal Clerk	19	.431	Mixed
Receptionist	16	.926	F-dom
Secretary	21	.968	F-dom
Switchboard Operator	18	.944	F-dom
Word Processor Operator	20	.967	F-dom
Craft and repair:			
Auto Mechanic	20	.008	M-dom
Carpenter	21	.067	M-dom
Electrician	21	.014	M-dom
Locksmith	19	.036	M-dom
Maintenance Mechanic	20	.023	M-dom
Maintenance Painter	21	.041	M-dom
Mason	13	.088	M-dom
Plumber	20	.012	M-dom

NOTE.—M-dom = male-dominated; F-dom = female-dominated.

References

Beller, Andrea H. "Occupational Segregation by Sex: Determinants and Changes." *Journal of Human Resources* 17 (Summer 1982): 371–92.

Bergman, Barbara R. "Does the Market for Women's Labor Need Fixing?" *Journal of Economic Perspectives* 3 (Winter 1989): 43–60.

Bielby, William T., and Baron, James N. "A Woman's Place Is with Other Women: Sex Segregation within Organizations." In *Sex Segregation in the Work Place: Trends, Explanations, Remedies,* edited by Barbara F. Reskin, pp. 27–55. Washington, DC: National Academy Press, 1984.

Blank, Rebecca M. "An Analysis of Workers' Choice between Employment in the Public and Private Sectors." *Industrial and Labor Relations Review* 38 (January 1985): 211–24.

Blau, Francine D. *Equal Pay in the Office.* Lexington, MA: Lexington Books, 1977.

England, Paula. "The Failure of Human Capital Theory to Explain Occupational Sex Segregation." *Journal of Human Resources* 17 (Summer 1982): 358–70.

Filer, Randall. "Sexual Differences in Earnings: The Role of Individual Personalities and Tastes." *Journal of Human Resources* 18 (Winter 1983): 87–99.

Freeman, Richard B. "Supply Elasticities for Educated Labor." In *Economics of Education: Research and Studies,* edited by George Psacharapolous, pp. 244–48. Oxford: Pergamon Press, 1987. (*a*)

———. "How Do Public Sector Wages and Employment Respond to Economic Conditions?" In *Public Sector Payrolls,* edited by David A. Wise, pp. 183–213. Chicago: University of Chicago Press, 1987. (*b*)

Groshen, Erica L. "The Structure of the Female/Male Wage Differential." *Journal of Human Resources* 26 (Summer 1991): 457–72.

Iowa Department of Personnel. "Minimum Qualifications Guide." Loose leaf. Des Moines, IA: Merit Employment, 1986.

Kim, Marlene. "Gender Bias in Compensation Structures: A Case Study of Its Historical Basis and Persistence." *Journal of Social Issues* 45 (Winter 1989): 39–50.

Krueger, Alan B. "Are Public Sector Workers Paid More than Alternative Wage? Evidence from Longitudinal Data and Job Queues." In *When Public Sector Workers Unionize,* edited by Richard B. Freeman and Casey Ichniowski, pp. 217–42. Chicago: University of Chicago Press, 1988.

McDowell, John M. "Obsolescence of Knowledge and Career Publication Profiles: Some Evidence of Differences among Fields in Costs of Interrupted Careers." *American Economic Review* 72 (September 1982): 752–68.

Orazem, Peter F., and Mattila, J. Peter. "The Implementation Process of Comparable Worth: Winners and Losers." *Journal of Political Economy* 98 (February 1990): 134–52.

———. "Human Capital, Uncertain Wage Distributions, and Occupational and Educational Choices." *International Economic Review* 32 (February 1991): 103–22.

Polachek, Solomon W. "Occupational Self-Selection: A Human Capital Approach to Sex Differences in Occupational Structure." *Review of Economics and Statistics* 58 (February 1981): 60–69.

Quinn, Joseph F. "Pension Wealth of Government and Private Sector Workers." *American Economic Review* 72 (May 1982): 283–87.

Sandell, Steven H., and Shapiro, David. "Work Expectations, Human Capital Accumulation, and the Wages of Young Women." *Journal of Human Resources* 15 (Summer 1980): 335–53.

Smith, Sharon P. *Equal Pay in the Public Sector: Facts or Fantasy.* Princeton, NJ: Princeton University Press, 1977.

Sorenson, Elaine. "Effect of Comparable Worth Policies on Earnings." *Industrial Relations* 26 (Fall 1987): 227–39.

———. "The Crowding Hypothesis and Comparable Worth." *Journal of Human Resources* 25 (Winter 1990): 55–89.

U.S. Bureau of Labor Statistics. "Employer Costs for Employee Compensation—March." *News.* Various issues.

———. *Employment and Earnings,* vol. 36 (January 1989).

———. *1980 Census of Population: Detailed Population Characteristics.* Washington, DC: U.S. Government Printing Office, 1983.

U.S. Bureau of the Census. Current Population Surveys. Tapes. Washington, DC: U.S. Bureau of the Census, various years.

U.S. Department of Labor. *Dictionary of Occupational Titles.* 4th ed. Washington, DC: U.S. Government Printing Office, 1977.

White, Halbert. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity." *Econometrica* 48 (May 1980): 817–38.