Title
Occupational segregation and the devaluation of women's work across US labor markets

Journal
Social Forces, 81(3)

ISSN
0037-7732

Authors
Cohen, Philip N
Huffman, M L

Publication Date
2003-03-01

Peer reviewed

Occupational Segregation and the Devaluation of Women's Work across U.S. Labor Markets*

Philip N. Cohen, *University of California at Irvine*
Matt L. Huffman, *University of California at Irvine*

Abstract

Previous research on the devaluation of women's work has investigated whether the net effect of gender composition varies across jobs and organizational settings. We extend that research by using hierarchical linear models that combine data from a random sample of U.S. work establishments with metropolitan-area data to explore whether macro-level gender inequality also influences the tendency to devalue women's work roles. Thus, we offer the first attempt to examine processes that lead to organizational gender inequality in local labor market contexts. Specifically, we hypothesize that gender devaluation will be strongest in highly gender-segregated labor markets. One reason for this may be that in segregated markets, men are in a stronger position to benefit from devaluation while women are less able to resist it. The results strongly support this hypothesis: Higher levels of occupational segregation at the labor market level are associated with a significantly increased tendency to devalue women's work roles. This finding is not explained by a diverse set of controls at both the establishment and local labor market level. Our findings highlight an additional way that gender segregation intensifies labor market inequality.

* *The authors contributed equally to the preparation of this manuscript; we list our names alphabetically. An earlier version of this article was presented at the 2001 meetings of the American Sociological Association. Some of the data used in this analysis are derived from sensitive data files of the 1991 National Organizations Study (NOS), obtained under special contractual arrangements designed to protect the anonymity of respondents. These data are* not *available from the authors. People interested in obtaining NOS sensitive data files should contact David Knoke, Department of Sociology, University of Minnesota, Minneapolis, MN 55455. E-mail: knoke001@maroon.tc.umn.edu. We thank Reeve Vanneman and his colleagues for sharing their Metropolitan Area data file with us. We also thank Michelle Budig, Marta Elvira, Julie Kmec, Barbara Reskin, Judy Stepan-Norris, Donald Tomaskovic-Devey, and anonymous* Social Forces *reviewers for their very helpful comments and suggestions. The authors are responsible for any remaining errors or omissions. Direct correspondence to Philip N. Cohen, Department of Sociology, University of California at Irvine, Irvine, CA 92697-5100. E-mail: cohenp@uci.edu.*

Social Forces, March 2003, 81(3):881-907

Earnings inequality by gender may result from discrimination both in the allocation of workers across job categories and in how female-dominated jobs are rewarded relative to male-dominated jobs (Petersen & Morgan 1995). The first form operates by blocking *access* to well-paying jobs and occupations through social closure processes (Tilly 1998); the second functions through the differential *valuation* of jobs and occupations with varying gender composition (Tomaskovic-Devey 1993b). Thus, the gender gap in pay is a function of two distinct sources: the differential distribution of women and men across jobs and occupations that vary with respect to pay, and within-job pay differences. Because it is uncommon to find women and men sharing the same job title in the same work establishment (Bielby & Baron 1984, 1986; Groshen 1991; Petersen & Morgan 1995; Tomaskovic-Devey 1993b), the first source — sorting workers into jobs — has been found to contribute the most to the pay gap.[1] However, explaining the mechanisms by which the determination of rewards operates is important for understanding how gender inequality in the labor market is constructed and perpetuated.

In this article, we reconsider the idea that the gender of the person who performs a particular line of work is an important determinant of how that work is rewarded, extending the large body of literature on gender-based ascription in the evaluation of job worth. By focusing on organizations in their labor market contexts, we directly address the effects of gender segregation at the labor market level. This enables us to unite research focusing on the impact of job and occupation-level gender segregation (Baron & Newman 1989; England et al. 1988; England, Reid & Kilbourne 1996; Huffman & Velasco 1997; Reid 1998) with literature highlighting the effect of macro-level, labor-market processes on earnings inequality (e.g., Beggs, Villemez & Arnold 1997; Cassirer 1996; Cohen 1998; Cotter et al. 1997; Jacobs & Blair-Loy 1996). We bring together these two bodies of research by examining whether status composition processes — those through which jobs performed by women accrue an economic penalty relative to those jobs held by men (see Tomaskovic-Devey 1993a) — are particularly strong when those jobs are embedded in segregated labor markets. We argue that gender devaluation should be strongest in highly gender-segregated labor markets. One reason for this may be that in segregated markets men occupy superior jobs in terms of pay and authority relative to women. As a result, men may be in a stronger position to benefit from devaluation, while women are less able to resist it. Therefore, our main hypothesis is that the negative association between female representation in a job and average wages will be especially pronounced when gender segregation in the broader labor market is high.

Gender-Based Devaluation: What People Do, Where They Work, or Both?

Comparable worth policies — predicated on the idea that women earn less because they work in female-dominated jobs and occupations — attempt to correct these between-job pay differences by adjusting pay scales so that both "male" and "female" jobs are compensated according to skill requirements (Acker 1989; England 1992; England & Dunn 1988). This policy approach is fueled by empirical investigations of the wage effect of gender composition. One line of research (England 1992; England et al. 1994; Parcel 1989) uses U.S. Census occupations as the units of analysis, with average earnings in the occupation as the dependent variable. These studies often include detailed measures of average educational requirements and other occupational demands taken from the *Dictionary of Occupational Titles* (U.S. Department of Labor 1977). Similar studies (Johnson & Solon 1986; Reid 1998; Tam 1997) use individual-level wages as the dependent variable and append occupational characteristics (including the occupation's gender composition) as contextual variables. These analyses often estimate the interaction between gender and demographic composition by fitting models separately for women and men; some (Macpherson & Hirsch 1995) use models with fixed effects in order to account for unmeasured gender differences in unchanging attributes of individuals relevant for pay determination (England et al. 1988).

This line of research consistently finds a significant net wage penalty associated with female-dominated occupations. For example, the findings of England and her colleagues (1988) include wage differences between male- and female-dominated occupations ranging from 8% to 11%, depending on the model specification. Kilbourne and her colleagues (1994) and Macpherson and Hirsch (1995) report this difference to be approximately 5%, while England, Reid, and Kilbourne (1996) find a wage differential ranging from 7% to 19% between male- and female-dominated occupation-industry cells. On the other hand, Tam (1997) argues that the statistical models and measures used in previous studies of the wage effect of gender composition bias the results to favor the devaluation hypothesis. Using a model that omits many commonly used measures, Tam (1997) reports no significant negative effect of occupation gender composition on wages.[2]

A second group of researchers (Baron & Newman 1990; Huffman & Velasco 1997; Petersen & Morgan 1995; Tomaskovic-Devey 1993a, 1995) uses job-level measures of female representation rather than the broader occupation-level measures. This allows them to examine the organizational contexts of reward allocation, as well as how the undervaluation of work roles performed by women or minorities varies by work setting and characteristics of the organizational environment.[3] Analyzing a sample of U.S. work establishments, Huffman and Velasco (1997) report a large job-level penalty associated with female representation that was not explained by, and did not vary systematically with, various measures of organizational structure. That is, not only did the gender composition effect remain large in the presence of controls for organizational policies, practices, and skill demands, but gender composition also did not interact with these measures. This net effect of gender composition suggests a remarkable stability in the degree of job-level devaluation in U.S. work establishments. Across jobs in diverse work establishments, as the level of female representation increases, the wages of both male and female employees decrease markedly.

This body of work also has indicated that as the level of classification becomes finer, the wage effect of gender composition becomes stronger. For example, in comparing the relative effect of occupational and job gender composition, Huffman, Velasco, and Bielby (1996) find that the typical earnings in an all-female job are about 50% lower than in a comparable all-male job. In another job-level study, Baron and Newman (1990) report that the difference in starting wages between male- and female-dominated civil service jobs is approximately 30%. Of course, the fact that the job-level gender composition effect on wages is much larger than that found in occupation-level studies is explained, at least in part, by the fact that measures of occupational segregation understate the degree of segregation occurring at the job level (Bielby & Baron 1986; Tomaskovic-Devey 1995). We agree with Tomaskovic-Devey (1995:24) that observed occupation-level effects are largely a result of more proximate processes operating at the level of the job (also see Reskin 1993).

In sum, empirical tests of the devaluation hypothesis have been nearly unanimous in their finding that female representation conditions the reward levels, even after controlling for job-related skills and educational requirements. Additionally, gender composition effects are explained neither by labor market conditions such as job growth (Baron & Newman 1990:156) nor by differences across jobs or occupations in working conditions that might constitute "compensating differentials" (England 1992; Huffman & Velasco 1997; Jacobs & Steinberg 1995; Reid 1998).[4] This finding characterizes nearly all studies conducted at the occupation level, as well as those performed at the more detailed job level.

What processes might explain the negative effect of female representation on reward levels? Kilbourne and her colleagues (1994) offer two potential mechanisms. First, employers err cognitively by not seeing women's contribution to the profitability of their organizations. Second, skills typically associated with women's work (such as nurturance) are underrewarded (see England et al. 1994). Both experimental and survey research on how work is valued support this position by showing that both sexes attribute more value to work performed by men than by women (see Bose & Rossi 1983; Deaux 1985; Major, McFarlin & Gagnon 1984; McArthur 1985). Others suggest that over time, occupations become defined largely by stereotypes based on the sex of the occupation's "typical" incumbent (Jacobs & Blair-Loy 1996; Milkman 1987). Once established, institutional inertia keeps both gender-based interoccupational and interjob wage differences in place and resistant to change over time (Baron 1991; Hannan & Freeman 1984; Kilbourne et al. 1994). The persistence of either of these mechanisms may also follow from gender dynamics within organizations, such that incumbents in female-dominated jobs have less decision-making power regarding pay and authority (Reskin, McBrier & Kmec 1999; Tomaskovic-Devey 1993b). Regardless of its source, we refer to the outcome as "devaluation" — lower wages in female-dominated jobs. We argue that this devaluation of women's work might not be invariant across diverse labor market contexts. Although most of the research on spatial variation in inequality focuses on race (see Beggs, Villemez & Arnold 1997; Cohen 2001; Tienda & Lii 1987), we address how variation in the gender devaluation effect may depend on local labor market factors.

A Missing Link: The Local Labor Market Context of Work Establishments

Although gender inequality is systemic throughout the U.S. economy, there is also pronounced variation across local labor markets (Cotter, Hermsen & Vanneman 1999). Gender devaluation also may be partly a function of local conditions, as the culture and power structure of the local labor market will have affected the composition of decision makers and the context in which they make their decisions over time. Beggs (1995) posits that the level of support for equality of opportunity in firms' local institutional environment influences the degree of racial and gender inequality. Additionally, Cotter and his colleagues (1997) argue that, in markets with more gender equality in occupational allocation, "changes in market pressures, normative expectations, and managerial power induced by occupational integration" may produce better jobs and higher earnings for women across the board (715–16). Their analysis shows that occupational gender segregation in local labor markets negatively affects women's earnings even outside of segregated occupations.

A stiffer local regime of gender devaluation could produce this effect. If gender devaluation were reproduced locally in this manner, it would represent an important mechanism linking the gendered practices of everyday life (West & Zimmerman 1987) with the larger structures of gender inequality in the economy (Goldin 1990). On this point, Nelson and Bridges (1999) cite research showing that historically, employment patterns in organizations often reflect status differences in the local community; they note that between-job wage inequality may result from the organizational reproduction of male advantage. Moreover, they assert that the devaluation of work done primarily by women "may have its most pronounced influence on pay disparities in *interaction with the culture and structure of employing organizations*" (95).

Previous research also shows that employers respond to changes in the legal environment with regard to gender inequality issues (Kelly & Dobbin 1999) and to workplace due-process procedures more generally (Edelman 1990). Some of this is observed at the local or state level (Guthrie & Roth 1999b). In fact, despite the effect of highly visible national trends and related campaigns (Guthrie & Roth 1999a), much of the pressure against gender inequality is generated locally. In the case of the law, for example, although the federal government provides national standards regarding employment discrimination, local laws allow cities to set higher standards (Gold 1993; Gutman 1993).[5]

These considerations suggest possible shortcomings of existing studies of the devaluation hypothesis. In the determination of wages, establishment-level characteristics partly may be a proxy for labor market attributes. This would be the case if the distribution of types of establishments or patterns of establishment practices were not uniform across labor markets. For example, Reskin and McBrier (2000) show that firms' personnel practices are associated with gender inequality in managerial hiring. Such practices may reflect local variation in acceptable standards, in which case the establishment variable would reflect local labor market characteristics. Given the importance of establishment practices, the distribution of these practices across labor markets may be an important component of how gender inequality is maintained — a pattern that would only emerge in an analysis that accounts for both levels. Previous studies of work-based devaluation are insensitive to potential interactions between the local context and the demographic composition of jobs, despite the fact that there is ample reason to ask whether gender-based devaluation is positively correlated with levels of occupational gender segregation at the local labor market level. Our analysis, in contrast to work that implies that devaluation is a uniform cultural process (England 1992), explores spatial variation in the gender devaluation effect across labor markets.

Data, Measures, and Analyses

The National Organizations Study

Our primary data source is the National Organizations Study (NOS), a random sample of 688 U.S. work establishments.[6] The 1991 General Social Survey (GSS) included a topical module on work organizations. That survey was used to draw the NOS sample, generating a random sample of U.S. work establishments with probability of inclusion proportionate to workplace size. Employed GSS respondents were asked to supply the contact information of both their employer and their spouse's employer. Subsequently, interviews with "knowledgeable organizational informants" at these organizations were conducted by telephone and questionnaires were mailed to them. These informants, who were heads of personnel departments or people responsible for making hiring decisions, were queried about the establishment's demographic composition, environment, and structure, as well as characteristics of specific job titles.[7]

Many of the NOS questions targeted the *core job* in the establishment, which is the job title of those employees most directly associated with the production of the firm's main service or product (e.g., production workers in a factory). The informants reported the "annual earnings of most core employees" in each NOS establishment, which, when logged, serves as the dependent variable in our multivariate models. The other NOS establishment-level variables we use consist of those variables that have been linked to gender inequality in previous research (see Baron, Mittman & Newman 1991; Huffman & Velasco 1997; Reskin & McBrier 2000; Tomaskovic-Devey 1993b). The primary independent variable of interest is the *proportion female in the core job title.*

We control for other aspects of the establishments' demographic composition by including variables measuring *proportion female in management* and *proportion nonwhite* (among full-time workers). We also include continuous measures of the establishments' *age* and *size* (both are logged, and size is equal to the sum of full- and part-time workers). Dichotomous variables are used to index whether the establishment is *part of a larger organization* (1 = yes), and whether it has a *professional/managerial core job* (1 = yes). Three dummy variables capture differences across four establishment *types* (corporate, government, nonprofit, and entrepreneurial). Nonprofit is the reference category.

Formalization of the establishments' rules and procedures was measured with an additive index composed of eight dichotomous indicators. These indicate the presence of these documents at the establishment: a rules and procedures manual, written job descriptions, written record of job performance, employment contracts, personnel evaluation documents, hire-fire procedure documents, documents describing safety-hygiene practices, and fringe benefits documents. The formalization index takes a value equal to the number of documents out of eight present. The extent to which a NOS establishment is subject to *government regulation* is based on the question, "How much of your establishment's operations are regulated by government agencies?" The possible answers ranged from "almost completely" to "none at all." We reverse-coded the original item so that high values indicate more regulation. The degree to which each establishment relies on an *internal labor market* to promote currently employed workers is measured with an additive index made up of three NOS items (see Marsden, Cook & Knoke 1994). The first is whether the establishment fills vacancies in the core job title with current employees (1 = yes; 0 = no). The second item asks whether different levels of the core job title exist (1 = yes; 0 = no). The final item taps the frequency of promotion above the core job title (3 = very often; 2 = often; 1 = not very often). The internal labor market index takes a value equal to the mean computed across the three indicators.

We also use 10 dummy variables to represent the 11 *industry* categories represented in the Office of Management and Budget's Standard Industrial Classification. The 11 categories are: agriculture, construction, nondurable manufacturing, durable manufacturing, transportation, wholesale trade, retail, F.I.R.E. (finance, insurance and real estate), business/repair and personal services, professional and related services, and public administration. Nondurable manufacturing is the omitted category in all multivariate models.

If there are differences in required skills, education, and working conditions across jobs, and these differences are correlated with both gender composition and wages, then the association between proportion female and reward levels could be spurious. Ideally, we would control for these characteristics at the level of the individual worker (education, skills) or job (working conditions, average training required, level of authority). Unfortunately, our data do not include individual-level attributes. We can employ some proxies, however, for job-level skills and requirements, from the *Occupational Measures from the Dictionary of Occupational Titles (DOT) for 1980 Census Detailed Occupations* data set (England & Kilbourne 1988). These are common in studies of gender and race composition effects on remuneration (see England 1992; Huffman & Velasco 1997; Tam 1997). The DOT measures are averages for workers in each 3-digit occupational category and are merged with the NOS data using the census 3-digit occupational code associated with the core job title of each establishment.

Specifically, we use three DOT variables as controls. The first is *specific vocational preparation* (SVP), which taps the amount of training time needed to learn the techniques and obtain the information necessary for average performance on the job. High values of this scale represent a longer period of time required to acquire these skills. SVP can be thought of as a measure of specific human capital (Tam 1997). We also use *general educational development* (GED), which measures "the typical requirement of the occupation for schooling that is not vocationally specific" (England, Hermsen & Cotter 2000:1742). GED is measured along a 0–6 point scale, with high values denoting those occupations requiring relatively high levels of general educational development. SVP and GED are positively correlated (r = .89, p < .01). Finally, we use a scale of *physical demands* from the DOT data set. The value of this scale is the average computed across five individual physical demands factors included in the DOT: climbing, stooping, reaching, talking, and seeing. High values of this variable indicate greater levels of physical demands.[8]

LOCAL LABOR MARKETS

Our second source of data is a metropolitan area (MA) labor market file constructed by Cotter and his colleagues (1997), mostly from the Census Bureau's STF3C file from the 1990 census. Each NOS establishment was mapped into a county by ZIP code; each county is part of only one MA in the file. Because ZIP codes can span counties (and therefore MAs), a particular NOS establishment can have one of the following four relationships to MAs. It can be located (1) in a single MA, (2) partially in an MA and partially not in any MA, (3) simultaneously in two different MAs, or (4) outside all MAs. Establishments found to be located partially in an MA and partially not in any MA were coded as part of that MA. Establishments that have a ZIP code straddling two MAs were assigned to the MA with the larger percentage of the ZIP code population within it. Establishments not located in an MA were assigned an MA if they were located within 75 miles of the central city of the nearest MA. If a rural establishment fell within this limit, it was assigned to the MA with the shortest distance between the establishment's ZIP code and the MA's central city. Establishments located more than 75 miles from the central city of any MA were excluded from the analysis. This yielded 475 establishments nested in 80 different MAs for use in the analyses.

The primary MA-level independent variable is *occupational integration*. A standard measure of occupational segregation reflects the proportion of women or men who would have to switch occupations in order to equalize the distribution of women and men across occupations. We subtract this measure from 1, and refer to it as occupational integration, with higher values indicating greater gender integration.[9] For the 80 MAs represented in the NOS sample, occupational integration ranges from .40 to .56. Examples of highly segregated labor markets include Lafayette, Louisiana, and Johnstown, Pennsylvania. Among the most integrated labor markets are Madison, Wisconsin, and Bryan–College Station, Texas. Although the NOS sample includes only 80 of the 262 possible U.S. MAs, the range computed across the 80 labor markets included in the NOS sample is virtually identical to the range based on the 262 U.S. MAs.[10]

Because of the relatively small number of metropolitan areas, we include only a limited number of MA-level control variables that have been shown to affect gender inequality in earnings (Cotter et al. 1997). These include variables reflecting basic economic structure and historical conditions: the percentage of the labor force employed in *durable-goods manufacturing* and dummy variables indicating each of the four census *regions* of the country (Northeast, North Central, South, and West). Others reflect local economic conditions: the net percentage change in the population resulting from 1985–90 *internal migration* (a proxy for long-term regional economic vitality), the *unemployment rate* (for short-term vitality), and the level of *male earnings inequality* (the

TABLE 1: Descriptive Statistics and Bivariate Correlation Coefficients for Establishment-Level Variables Used in the Analysis

	1	2	3	4	5	6	7	8	9
1. Earnings of core job title (ln)	1.0	—	—	—	—	—	—	—	—
2. Proportion female in core job title	–.25*	1.0	—	—	—	—	—	—	—
3. Proportion female management	–.15*	.57*	1.0	—	—	—	—	—	—
4. Establishment size (ln)	.35*	.05	–.09*	1.0	—	—	—	—	—
5. Establishment age (ln)	.32*	–.04	–.10*	.44*	1.0	—	—	—	—
6. Proportion nonwhite	–.06	.03	.03	.29*	.04	1.0	—	—	—
7. Part of a larger organization	.11*	.02	–.12*	.32*	.17*	.11*	1.0	—	—
8. Regulation	.25*	.05	–.03	.35*	.28*	.08	–.10*	1.0	—
9. Professional/managerial core job title	.38*	.21*	.21*	.19*	.22*	–.02	–.34*	.18*	
10. Institutionalization scale	.27*	.01*	.01	.45*	.25*	.05	.24*	.83*	.23*
11. Formalization	.32*	.07	–.01	.66*	.33*	.18*	.49*	.40*	.17*
12. Corporate establishment	–.06	–.13*	–.19*	.06	–.05	.02	.03	–.10*	–.34*
13. Government establishment	.13*	–.19*	–.20*	.13*	.26*	.02	.24*	.32*	.02
14. Entrepreneurial establishment	–.32*	.01	.15*	–.48*	–.38*	–.07	–.35*	–.36*	–.20*
15. Internal labor market	.28*	–.03	–.14*	.62*	.28*	.18*	.39*	.27*	.12*
16. Standard vocational preparation (SVP)	.43*	.01	.11*	.07	.20*	–.17*	–.04	.09	.69*
17. General educational development (GED)	.42*	.18*	.22*	.23*	.23*	–.15*	.02	.13*	
18. Physical demands scale	–.15*	–.26*	–.16*	–.35	–.16*	.06	–.06	–.15*	–.53*
Mean	9.94	.47	.37	4.18	3.10	.21	.54	3.37	.33
Standard deviation	.66	.40	.35	2.25	1.21	.26	.50	1.23	.47
Minimum	7.44	0	0	0	0	0	0	1.0	0
Maximum	12.0	1.0	1.0	10.7	5.32	1.0	1.0	5.0	1.0

(N = 475)

earnings Gini coefficient for full-time year-round employed men in the age range 25–54). Three variables measure the local demographic structure: the natural logarithm of the *size of the labor force*, the percentage of the population that is *black*, and the percentage that is *Latino*.

Finally, given its importance in explaining a wide range of outcomes related to gender stratification (Chafetz 1984; Cotter et al. 1998), we control for the

TABLE 1: Descriptive Statistics and Bivariate Correlation Coefficients for Establishment-Level Variables Used in the Analysis (Cont'd)

	10	11	12	13	14	15	16	17	18
1. Earnings of core job title (ln)	—	—	—	—	—	—	—	—	—
2. Proportion female in core job title	—	—	—	—	—	—	—	—	—
3. Proportion female management	—	—	—	—	—	—	—	—	—
4. Establishment size (ln)	—	—	—	—	—	—	—	—	—
5. Establishment age (ln)	—	—	—	—	—	—	—	—	—
6. Proportion nonwhite	—	—	—	—	—	—	—	—	—
7. Part of a larger organization	—	—	—	—	—	—	—	—	—
8. Regulation	—	—	—	—	—	—	—	—	—
9. Professional/managerial core job title	—	—	—	—	—	—	—	—	—
10. Institutionalization scale	1.0	—	—	—	—	—	—	—	—
11. Formalization	.49*	1.0	—	—	—	—	—	—	—
12. Corporate establishment	–.05	.06	1.0	—	—	—	—	—	—
13. Government establishment	.16*	.21*	–.35*	1.0	—	—	—	—	—
14. Entrepreneurial establishment	–.40*	–.53*	–.44*	–.21*	1.0	—	—	—	—
15. Internal labor market	.31*	.58*	.13*	.11*	–.41*	1.0	—	—	—
16. Standard vocational preparation (SVP)	.11*	.03	–.28*	–.01	–.09*	.06	1.0	—	—
17. General educational development (GED)	.19*	.10*	–.31*	–.01	–.16*	.08	.89*	1.0	—
18. Physical demands scale	–.04	–.04	.14*	.06	.11*	–.03	–.26*	–.52*	1.0
Mean	1.72	.75	.42	.14	.21	.74	5.38	3.85	1.66
Standard deviation	.49	.34	.49	.35	.41	.47	1.55	.94	.88
Minimum	.5	0	0	0	1	0	1.7	1.56	0
Maximum	2.5	1.0	1.0	1.0	1.0	1.67	8.5	6.0	3.90

(N = 475)

* p < .01 (two-tailed)

demand for female labor implied by the occupational structure. Specifically, this variable taps the share of the labor force that would be female if local occupations had the same percentage female found in the country as a whole. It measures the degree to which the occupational structure of the local labor market is tilted toward female-dominated occupations. These variables are consistent with those used in most previous research on labor market earnings

TABLE 2: Descriptive Statistics and Bivariate Correlation Coefficients for Metropolitan Area–Level Variables Used in the Analysis

	1	2	3	4	5	6
1. Gender integration	1.0	—	—	—	—	—
2. Total population (ln)	.16	1.0	—	—	—	—
3. Percentage black	–.25*	.12	1.0	—	—	—
4. Percentage Hispanic	.08	.33*	–.16	1.0	—	—
5. Durable good manufacturing	–.12	.11	–.24	–.13	1.0	—
6. Male earnings Gini	.15	.41*	.12	.64*	–.47*	1.0
7. Unemployment rate	–.47*	–.004	.07	.52*	.17	.10
8. Female labor demand	.48*	.08	–.01	–.004	–.44*	.17
9. Net migration	.27	–.02	.12	–.11	–.12	.08
10. Northeast	–.04	.30*	–.08	–.07	.05	–.06
11. North central	.002	–.11	–.16	–.25	.37*	–.38*
12. South	–.25	–.28*	.52*	.001	–.45*	.28*
13. West	.34*	.21	–.37*	.36*	.07	.16
Mean	.50	13.99	.12	.07	.11	.35
Standard deviation	.06	3.52	.23	.31	.12	.06
Minimum	.40	11.02	.001	.003	.03	.29
Maximum	.56	16.78	.46	.85	.26	.42

variation (see Fossett 1988; Grant & Parcel 1990), and almost all have been shown to affect gender inequality in labor markets (Cotter et al. 1997). Descriptive statistics and bivariate correlations for all variables used in the analysis are reported in Tables 1 and 2.

Statistical Model and Analysis Strategy

Our analysis employs hierarchical linear models (Bryk & Raudenbush 1992) that combine the NOS establishment-level variables with data from U.S. metropolitan areas.[11] Hierarchical modeling has become more popular as its advantages have been documented and as increased computing resources and new software have made its application more practical. According to Guo and Zhao (2000), hierarchical models are used for analyzing nested data structures because they correct biases in parameter estimates resulting from clustering in the data, and because they provide correct standard errors and significance tests when data are hierarchically structured.[12]

Implicitly, our model allows an establishment-level equation to be estimated separately in each of the 80 MAs. Each establishment-level coefficient may be treated as a dependent variable in the MA-level analysis. We are particularly interested in the MA-level determinants of the effect of proportion female in

TABLE 2: Descriptive Statistics and Bivariate Correlation Coefficients for Metropolitan Area-Level Variables Used in the Analysis (Cont'd)

	7	8	9	10	11	12	13
1. Gender integration	—	—	—	—	—	—	—
2. Total population (ln)	—	—	—	—	—	—	—
3. Percentage black	—	—	—	—	—	—	—
4. Percentage Hispanic	—	—	—	—	—	—	—
5. Durable good manufacturing	—	—	—	—	—	—	—
6. Male earnings Gini	—	—	—	—	—	—	—
7. Unemployment rate	1.0	—	—	—	—	—	—
8. Female labor demand	–.32*	1.0	—	—	—	—	—
9. Net migration	–.30*	–.11	1.0	—	—	—	—
10. Northeast	–.001	–.35	–.15	1.0	—	—	—
11. North central	–.12	–.008	–.24	–.27*	1.0	—	—
12. South	.08	–.17	.27	–.30*	–.49*	1.0	—
13. West	.04	–.11	.10	–.20	–.32*	–.36*	1.0
Mean	.06	.46	.001	.15	.31	.36	.19
Standard deviation	.04	.03	.09	.86	1.13	1.18	.97
Minimum	.03	.42	-.15	0	0	0	0
Maximum	.14	.50	.13	1.0	1.0	1.0	1.0

(N = 80)

Note: All statistics weighted by number of establishments per MA.

* p < .01 (two-tailed test)

the core job at the establishment level. This gender composition effect measures gender-based devaluation in each MA. Therefore, this model allows us to test whether occupational gender integration reduces the devaluation of jobs performed typically by women.

Specifically, our establishment-level model is described by

$$Y_{ij} = \beta_{00} + \beta_{1j}(ProportionFemale_{ij}) + \beta_{2j}(X_{2ij} - \overline{X}..) + ... + \beta_{kj}(X_{kij} - \overline{X}..) + r_{ij}$$

Where Y_{ij} equals the typical earnings (logged) of the core job title of establishment i in MA j, and β_0 is the level-1 intercept. β_{1j} is the regression coefficient associated with the core job proportion female in establishment i in MA j. X_{2ij} through X_{kij} is a set of $K - 1$ establishment-level control variables (each centered at its grand mean, $\overline{X}..$) and β_{kj} is a vector of regression coefficients associated with those control variables. Finally, r_{ij} is the level-1 error term, assumed to be normally distributed with mean zero and variance σ^2.

TABLE 3: Hierarchical Linear Regression Results for Logged Earnings of Core Job Title on Establishment-Level Characteristics

	Model 1 Coef.	Model 2 Coef.	Model 3 Coef.
Intercept	9.925***	10.119***	10.182**
Proportion female in core job	—	–.413***	–.529***
Establishment size (ln)	—	—	.049***
Establishment age (ln)	—	—	.035
Proportion female management	—	—	–.056
Proportion nonwhite	—	—	–.228*
Part of a larger organization	—	—	–.071
Regulation	—	—	.016
Professional/managerial core job	—	—	.162
Institutionalization scale	—	—	.018
Formalization	—	—	.316**
Corporate establishment	—	—	–.048
Government establishment	—	—	–.148
Entrepreneurial establishment	—	—	–.071
Internal labor market	—	—	.038
Standard vocational preparation (SVP)	—	—	.046
General educational development (GED)	—	—	.111
Physical demands scale	—	—	–.019
Level-1 R^2	—	.106	.450
Variance components			
Level-1 variance (σ^2)	.3932	.3670	.2298
Intercept (τ_{00})	.0397	.0198	.0082
Proportion female in core job (τ_{11})	0	.0256	.0441
Intraclass correlation coefficient (ρ)	.0917	.0512	.0345

Notes: Omitted categories are West region and nonprofit establishment. Industry (11 SIC categories) is also controlled; coefficients are not shown but are available upon request from the authors.

* p < .05 ** p < .01 *** p < .001 (two-tailed test)

The complete MA-level equation is given by:

$$\beta_{0j} = \gamma_{00} + \gamma_{01}(\mathit{Occupational Integration}_j) + \gamma_{02j}(W_{2j} - \overline{W}..) + ... + \gamma_{0sj}(W_{sj} - \overline{W}..) + u_{ij}$$

$$\beta_{1j} = \gamma_{10} + \gamma_{11}(\mathit{Occupational Integration}_j) + \gamma_{12j}(W_{2j} - \overline{W}..) + ... + \gamma_{1sj}(W_{sj} - \overline{W}..) + u_{ij}$$

$$\beta_{kj} = \gamma_k$$

where γ_{00} is the intercept for the MA-level model of typical earnings (logged) in the core job title, and γ_{01} is the effect of labor market occupational integration on β_{0j}. The MA-level intercept for the effect of core job proportion

TABLE 4: Hierarchical Linear Regression Results for Logged Earnings of Core Job Title on Establishment-Level and Local Labor Market Characteristics

	Model 1 Coef.	Model 2 Coef.
Intercept		
Intercept	10.076***	11.913***
Gender integration	.214	–3.562**
Percentage durable goods	—	–1.126
Northeast	—	.033
North central	—	–.157
South	—	–.173†
Unemployment rate	—	–5.892*
Male earnings Gini	—	–.647
Net internal migration	—	.450
Population size (ln)	—	.032
Percentage black	—	.530
Percentage Hispanic	—	.812*
Female labor demand	—	2.510
Proportion female in core job		
Intercept	–1.673	–4.655**
Gender integration	2.297	8.400**
Percentage durable goods	—	2.510
Northeast	—	.170
North central	—	.153
South	—	.189
Unemployment rate	—	6.702
Male earnings Gini	—	4.158
Net internal migration	—	–.098
Population size (ln)	—	–.025
Percentage black	—	–.543
Percentage Hispanic	—	–1.132
Female labor demand	—	–9.134
Variance components		
Level-1 variance (σ^2)	.2300	.2307
Intercept (τ_{00})	.0099	.0006
Proportion female in core job (τ_{11})	.0441	.0606
Intraclass correlation coefficient (ρ)	.0413	.0026

Notes: Omitted categories are West region and nonprofit establishment. Industry (11 SIC categories) and all establishment-level variables from Table 3 are also controlled; coefficients are not shown but are available upon request from the authors.

† $p < .10$ * $p < .05$ ** $p < .01$ *** $p < .001$ (two-tailed test)

female on typical earnings in the core job title is denoted by γ_{10}, and γ_{11} is the effect of labor market occupational integration on β_{1j}. W_{2j} through W_{sj} is a set of $S - 1$ establishment-level control variables (each centered at its grand mean, $\overline{W}_{..}$), while γ_{0sj} and γ_{1sj} are the vectors of regression coefficients associated with those control variables in each level-2 model. The effects of the level-1 control variables do not vary across MAs; therefore, γ_k represents the fixed effects β_k across all MAs. Finally, u_{0j} and u_{1j} are level-2 random effects, assumed to be uncorrelated and with means of zero. It is common to denote the variance of these level-2 error terms as t_{00} and t_{11}, respectively (see Bryk & Raudenbush 1992).[13] Among level-2 independent variables, only gender integration is not centered.

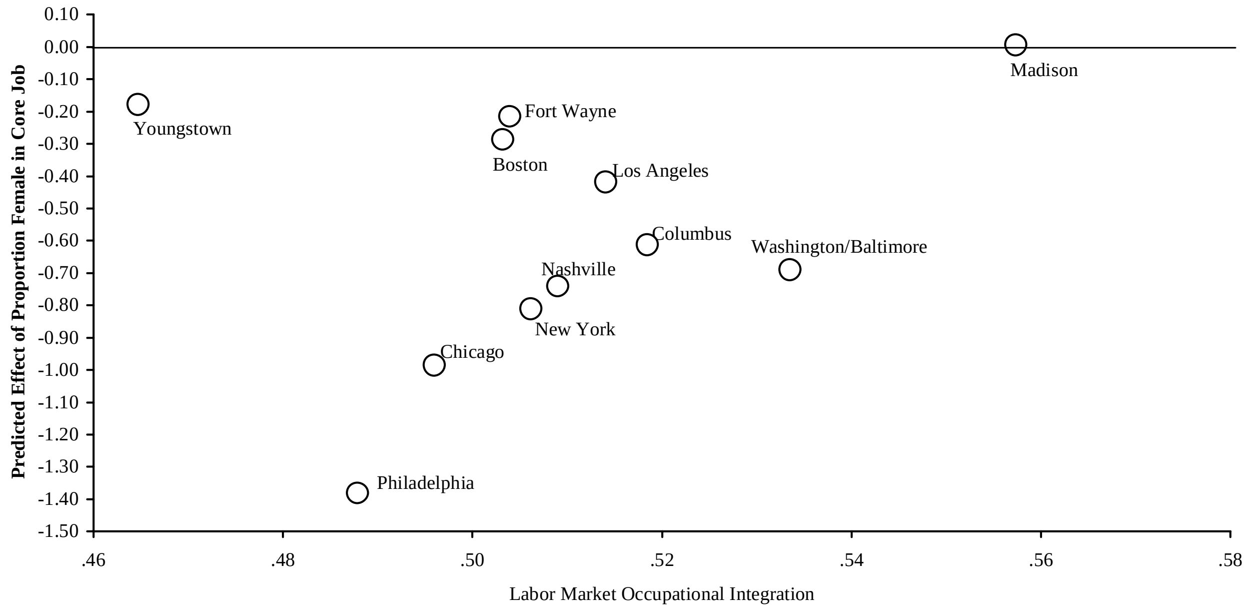
We begin by estimating a random intercept model, which includes no independent variables at either level. This model, which is equivalent to a one-way analysis of variance with random effects, expresses variation in the dependent variable (Y_{ij}) as the sum of the grand mean outcome in the population (γ_{00}), a level-1 random effect (r_{ij}), and a level-2 random effect (u_{0j}). This model is useful because it allows the total variability in the outcome variable to be decomposed into within- and between-group parts (Bryk & Raudenbush 1992; Snijders & Bosker 1999). Specifically, it can be shown that the variance of Y_{ij} is equal to $\tau_{00} + \sigma^2$, where τ_{00} equals the between-group variability, and σ^2 equals the within-group variation.[14]

Results

The results from the first set of hierarchical linear models appear in Table 3. Model 1 is a random intercept model; the intraclass correlation coefficient for this model indicates that about 9% of the total variability in wages is due to differences across MAs, while 91% is attributable to earnings differences across core jobs. This is not surprising, as one would expect the difference in earnings between lawyers and janitors in a particular MA to be much larger than the difference between lawyers in Boston and lawyers in Los Angeles, for example. In model 2, proportion female in the core job alone is shown to account for about 10% of the variability in between-job wages. Of the remaining variance, only about 5% is due to differences across MAs ($\rho = .0512$).

With no MA-level variables included, model 3 is similar to an ordinary least squares (OLS) regression, except that the intercept and the effect of core job proportion female are allowed to vary across MAs. Because all the establishment-level variables except core job proportion female are grand-mean-centered, the intercept in this model (10.182, or $26,423) is the predicted earnings for a typical establishment that has no women in its core job. The negative net effect of core job proportion female is large and statistically significant (–.529, $p < .001$). To give a sense of the magnitude of this effect, the

FIGURE 1: Gender Devaluation in Select Labor Markets, by Labor Market Occupational Integration



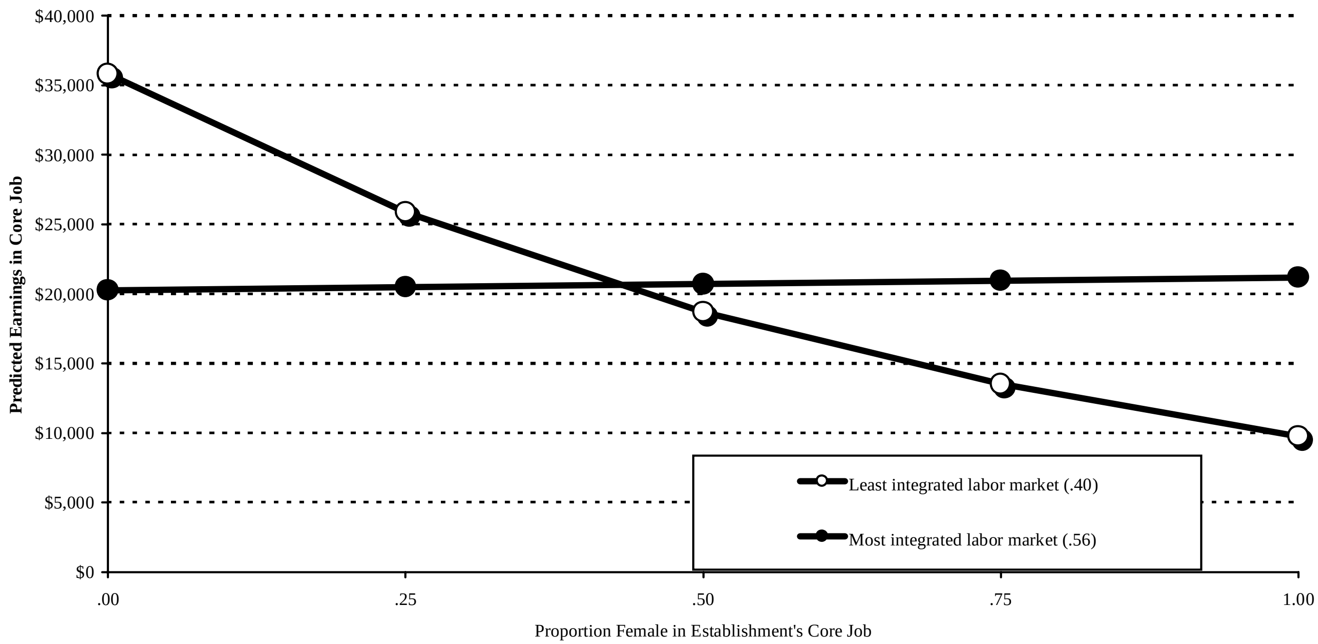
proportion female coefficient of –.529 means that as proportion female rises from 0 (all men) to 1.0 (all women), the earnings in the core job are predicted to decrease by 41% ($[e^{-.529} - 1] \times 100 = -41.1$). That is the gender devaluation effect.

The gender-composition effect we find is larger than that found in some previous research (for example, England et al. 1988; England, Reid & Kilbourne 1996; Kilbourne et al. 1994; Macpherson & Hirsch 1995). There are several probable reasons for this difference. First, as noted above, the wage effect of gender composition is stronger with finer levels of classification (Baron & Newman 1990; Huffman, Velasco & Bielby 1996), and our measures are at the establishment-specific job level. Second, we do not control for variables at the individual level. If women have, for example, lower average levels of education or are more likely to work part-time, and these traits are associated with lower wages — or if women's individual wage returns to these characteristics are different from men's — then our gender composition variable will capture some of the gender gap resulting from these differences as well. Finally, we note that some previous studies have collapsed jobs (or occupations) into male-dominated and female-dominated categories. As a result, the gender composition effect they find will be smaller than ours, which reflects predicted wage differences between jobs that are *all* male and those that are *all* female.

However, our primary objective in this paper is to analyze the difference in the level of gender devaluation across labor markets. Of most interest is how the gender devaluation effect is conditional on labor market segregation. Least squares estimates of the devaluation coefficient for each of the 11 labor markets (MAs) in our sample that have more than 10 establishments are shown in Figure 1 (obtained from model 3 of Table 3). This is the bivariate relationship between labor market integration and gender devaluation. The relationship suggests that gender devaluation is generally less pronounced in more integrated labor markets. In the large labor markets in our sample, then, the relationship between occupational gender segregation and gender devaluation is consistent with our main hypothesis: Gender segregation intensifies the negative effect of job percentage female.

In the subsequent models, shown in Table 4, we enter gender integration at the labor market level and test its effect on gender devaluation in the presence of the control variables at both levels. In these models, the core job proportion female and the labor market occupational integration variables are not centered. Therefore, the level-2 intercept (γ_{00}) represents the predicted earnings of an all-male job in a completely segregated local labor market, at the mean of the establishment and labor market control variables. And the intercept for the core job proportion female effect (γ_{10}) represents the effect of proportion female at 0 gender integration.[15] In the final model, only labor

FIGURE 2: Gender Devaluation in High- and Low-Gender Integration Labor Markets



market gender integration has a significant effect on the effect of job proportion female.[16]

The final model (model 2 of Table 4) predicts that in an establishment with only men in the core job, at the mean of the establishment controls, in a labor market at the mean of the controls but with complete occupational segregation, the log of the typical earnings will be 11.913 (or $149,194). Fixing job proportion female at 0 but allowing occupational integration to rise to complete integration (which is well beyond the range of the data), the expected log earnings plummets by 3.562, to 8.351 ($4,234). The intercept of –4.655 for core job proportion female reflects the gender devaluation associated with a rise from no women to all women in the core job, at 0 gender integration. As gender integration rises, it counteracts the effect of core job proportion female. At the hypothetical complete labor market integration, the effect of core job proportion female would actually be positive, 3.745 = –4.655 + 8.400.

Although core job proportion female ranges from 0 to 1, the actual range of occupational integration at the labor market level is only from .40 to .56, so these extreme values are never reached in our sample. To illustrate the predicted effects within the range of the data, Figure 2 shows the predicted earnings in the core job for establishments in the most and the least integrated labor markets, as proportion female in the core job rises from 0 to 1. The figure shows that the negative effect of proportion female on earnings is only apparent for establishments in less integrated labor markets. In the least integrated labor markets, predicted earnings fall from more than $35,000 to less than $10,000 as proportion female rises to 1. In the most integrated markets, the practical effect of job proportion female is predicted to be negligible.[17]

Thus, we find that the generalized gender devaluation effect found in previous research is most pronounced in labor markets that are segregated by gender. While previous research has documented gender devaluation at the organizational level, this tendency is very strongly linked to the local context of gender segregation. These results highlight another way that gender segregation disadvantages women — not only does segregation constitute the basis for devaluation by establishing gender-specific work roles, but it also *strengthens* the tendency for women's jobs to be paid less than comparable jobs performed by men.

Conclusion

A well-established body of empirical research demonstrates the close link between gender segregation (at various levels) and earnings inequality, showing, among other things, that female representation depresses wage rates. This research also addresses more specific questions about how this penalty accruing to work done primarily by women varies across diverse job- and establishment-level characteristics. Although there is no question that this work has added greatly to what we know about the dimensions of gender inequality, it does not address the important question of how the tendency to devalue women's work may be dependent on the level of gender segregation in the broader labor market. Thus, previous research has not asked whether devaluation is stronger or weaker in segregated labor market contexts.

To this end, we used a unique combination of data sets to explore how the tendency for women's work to be devalued is associated with occupational gender segregation at the labor market level. Specifically, we hypothesized that in labor markets with a relatively high level of gender segregation — a key indicator of gender inequality — men are in a stronger position to benefit from devaluation and women are less able to resist it. Likewise, labor markets with more gender segregation may have stronger systems of allocation and value assessment that have evolved in a more inequitable environment.

Our results support the hypothesis that higher levels of occupational segregation at the labor market level are associated with a strongly increased tendency for gender-based devaluation of women's work roles. The implications of our findings are substantial. Although devaluation certainly occurs at the establishment level (see Huffman & Velasco 1997; Tomaskovic-Devey 1993b), it takes place in a larger labor market context. This underscores the importance of macro-level conditions in the determination of inequality manifested at the establishment level and raises the possibility that in the absence of occupational segregation, at least some of the basis for gender-based devaluation would be substantially eroded. In previous research (Huffman & Velasco 1997), gender devaluation was found to be relatively constant across a range of organizational characteristics. However, devaluation does vary as a function of labor market characteristics, especially occupational gender segregation. This does not imply that organizational processes are irrelevant or that changing the practices of firms will not substantially affect those processes producing between-job wage inequality. Rather, it shows that their aggregate practices at the labor market level are associated with the level of gender devaluation.

Although our analysis addresses an important unanswered question regarding the organizational and labor market sources of gender inequality, several limitations should be noted. Most importantly, while other studies in this area use data on individuals but not their establishment-specific jobs (see England, Reid & Kilbourne 1996), we have data on jobs but not on individuals in those jobs. It is likely, therefore, that some of the gender-composition effect on wages that we find would be accounted for by individual-level factors, such as full-time versus part-time employment, level of education, age, marital status, and number of children. To narrow this gap, we control for the skill demands and training requirements for the occupation of each job, which captures some of this individual variation. Although this limitation is significant, in this analysis our primary concern is the variation in the devaluation effect across labor markets; we have no reason to believe that the omission of individual-level variables results in biased estimates of that variation.

Further, in the current analysis we cannot rule out the possibility that some part of the devaluation effect we have identified results from the lower wages women receive, generally, net of the control variables we include. Future research and data collection efforts need to include individual as well as job and labor market characteristics. In new research, we use U.S. Census microdata to examine individual workers nested within local occupation-industry cells (Cohen & Huffman 2002), which gives us a large sample of individuals and simulated jobs across labor markets (but sacrifices the job- and establishment-specific variables we have from the NOS). This results in estimates of the gender-composition effect on wages that account for individual worker characteristics. This also allows us to estimate *within-job* gender inequality — that is, how much men versus women earn in these jobs.

Finally, although our results clearly suggest that labor market segregation strengthens gender-based devaluation, our data do not allow us to precisely specify the mechanisms underlying this association. For example, Beggs (1995) suggests that both race and gender inequality reflect spatial variation in local institutional history, which we do not directly measure. Incorporating such variables into a multilevel framework would allow researchers to better understand the processes that create and sustain gender inequality in important labor market outcomes.

Notes

1. A well-established empirical finding is that the sex gap in pay decreases as occupational classifications become finer (Bielby & Baron 1984; Hultin & Szulkin 1999).

2. For a rebuttal of Tam's (1997) model specification and a reanalysis of his data, see England, Hermsen, and Cotter (2000).

3. Alternately, percentage female in individuals' occupation-industry cell has been used as a proxy for job-level gender composition (Budig 2002; England, Reid & Kilbourne 1996; Reid 1998).

4. As England, Reid, and Kilbourne (1996:513) note, some economists (Bergmann 1974, 1986) interpret the effect of sex composition as suggestive of "crowding" rather than devaluation. When access to some occupations is blocked, the oversupply of women in the remaining occupations results in lower pay relative to men. Although it is extremely difficult to differentiate crowding and devaluation empirically, Baron and Newman (1990) provide evidence that the net effect of demographic composition on job rewards reflects devaluation rather than crowding. See England (1992) and Jacobsen (1994) for detailed discussions of the sources and consequences of occupational crowding.

5. For example, see Bain (1985–86) and Daspit (1994). These local efforts also often involve fractious public debates that could heighten awareness of inequality and help generate pressure on employers. The mayor of San Francisco, Willie Brown, claims that the city "has shown that cities committed to the fight for justice and equality can prevail" and has sent a "message to other local governments around California and the nation" (Brown 1999). On attempts to publicize local levels of gender inequality in executive positions, see "Omaha Women Make Strides into Executive Ranks" (*Omaha World-Herald*, October 21, 2000); "The Culture Club: Diversity Efforts Make Some Inroads" (*Kansas City Star*, Sept. 19, 2000); and "Minorities Making Little Headway in Senior Management Positions" (*Milwaukee Journal Sentinel*, June 20, 2000). For efforts with regard to blue-collar workers, see "O.C. Far Below Average in Hiring Female Police" (*Los Angeles Times*, February 20, 2000); and, "MCLU: Too few women, people of color in volunteer fire departments" (*Star Tribune* [Minneapolis], Dec. 16, 1993).

6. The original NOS data set comprised 727 work establishments. However, 39 were duplicates.

7. Additional details regarding the NOS survey design can be found in Kalleberg et al. (1994) and Spaeth and O'Rourke (1996).

8. Tam (1997) argues that estimates of gender composition effects will be biased in favor of the devaluation hypothesis when both SVP and GED are included in the same model (but see England, Hermsen & Cotter 2000). Tam (1997), using only SVP and industry controls, finds no between-occupation gender composition effect. We replicated our level-1 model without GED but with our other controls — and again using just SVP and industry controls — and in all cases found a strong and significant ($p < .001$) negative effect of proportion female on earnings in the core job. These results are available from the authors upon request.

9. Rather than using the standard dissimilarity index (see Duncan & Duncan 1955), which is the most commonly used segregation measure, we use an adjusted dissimilarity index, taken from Cotter and colleagues (1997:730). It yields the proportion of women who would have to change occupations so that the observed number of women was no larger or smaller than chance would predict. This measure has the same interpretation as the unadjusted dissimilarity index, but it uses a random distribution as the standard rather than absolute equality.

10. More than 95% of the MAs nationally fall within the range of occupational integration computed across the 80 NOS MAs.

11. We estimated our models with the HLM software package, version 5.04. Estimates are derived through maximum likelihood estimation.

12. For example, two jobs in the same labor market will have identical values on all labor market variables, yielding downwardly biased standard errors.

13. The estimated variance components τ_{00} and σ^2 can be used to form the intraclass correlation coefficient (ρ), which yields the proportion of the variance in the outcome variable that exists between the level-2 units. It is computed as $\rho = \tau_{00} / (\tau_{00} + \sigma^2)$.

14. We also report a level-1 R^2 measure defined by Snijders and Bosker (1994). This measure is equal to the proportional reduction in the quantity $\tau_{00} + \sigma^2$ due to inclusion of predictors in the model. Following Snijders and Bosker (1994), we compute R^2 by comparing the quantity $\tau_{00} + \sigma^2$ of the random intercept model with the same quantity in a model that includes independent variables. Thus, $R^2 = 1– [\tau_{00} + \sigma^2$ for a model that includes predictors / $\tau_{00} + \sigma^2$ for a model with no predictors]. For example, for model 2 of Table 3, $R^2 = 1– [(.0198 + .3670) / (.0397 + .3932)] = .106$.

15. Although the coefficients for gender integration are quite different in the two models in Table 4, this is an artifact of the variable centering. To see the predicted effect of proportion female in the average MA, which has a gender integration of .50, we add one-half of the gender integration effect to the proportion female intercept. This yields a proportion female effect of –.525 in model 1 (–1.673 + [.50 × 2.297]), and –.455 in model 2 (–4.655 + [.50 × 8.400]).

16. In semi-reduced form models not shown, several of the control variables show significant effects, including those for durable goods manufacturing and female labor demand. The positive effect of durable goods manufacturing could help explain the position of Youngstown, Ohio, in Figure 1.

17. One reviewer suggested that our devaluation effect might reflect the greater propensity of female workers to work part-time. The NOS data do not include a measure of the proportion of core job workers who work part-time, but they do include the number of full- and part-time workers in the entire establishment (although this measure is missing in 110 of our 475 establishments). We use these variables to construct a variable indicating the proportion of workers who are part-time (imputing the mean score and adding a dummy variable for missing cases) to test the possibility that part-time workers are confounding our gender devaluation effect. In the final model, proportion part-time does have a significant negative effect, but the inclusion of this variable does not substantially change the rest of the model. The net effect of core job proportion female (at .50 occupational integration) is reduced from .455 to .434, but the effect of gender integration remains positive and significant. This is perhaps not surprising since 75% of our establishments have fewer than 29% part-time workers.

References

Acker, Joan. 1989. *Doing Comparable Worth: Gender, Class, and Pay Equity*. Temple University Press.

Bain, Adam J. 1985–86. "In Re Associated Sign & Post, Inc.: The Affirmative Action Obligations of Government Contractors in Indiana." *Indiana Law Journal* 61 Ind. L.J. 793.

Baron, James N. 1991. "Organizational Evidence of Ascription in Labor Markets." Pp. 113–43 in *New Approaches to Economic and Social Analyses of Discrimination*, edited by Richard R. Cornwall and Phanindra V. Wunnava. Praeger.

Baron, James N., Brian S. Mittman, and Andrew E. Newman. 1991. "Targets of Opportunity: Organizational and Environmental Determinants of Gender Integration within the California Civil Service, 1979–1985." *American Journal of Sociology* 96:1362–1401.

Baron, James N., and Andrew E. Newman. 1989. "Pay the Man: Effects of Demographic Composition on Prescribed Pay Rates in the California Civil Service." Pp. 107–30 in *Pay*

Equity: Empirical Inquiries, edited by Robert T. Michael, Heidi I. Hartmann, and Brigid O'Farrell. National Academy Press.

———. 1990. "For What It's Worth: Organizations, Occupations, and the Value of Work Done by Women and Nonwhites." *American Sociological Review* 55:155–75.

Beggs, John J. 1995. "The Institutional Environment: Implications for Race and Gender Inequality in the U.S. Labor Market." *American Sociological Review* 60:612–33.

Beggs, John J., Wayne J. Villemez, and Ruth Arnold. 1997. "Black Population Concentration and Black-White Inequality: Expanding the Consideration of Place and Space Effects." *Social Forces* 76:65–91.

Bergmann, Barbara. 1974. "Occupational Segregation, Wages, and Profits: When Employers Discriminate." *Eastern Economic Journal* 1:103–10.

———. 1986. *The Economic Emergence of Women*. Basic Books.

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

———. 1986. "Men and Women and Work: Sex Segregation and Statistical Discrimination." *American Journal of Sociology* 91:759–99.

Bose, Christine E., and Peter H. Rossi. 1983. "Gender and Jobs: Prestige Standing of Occupations As Affected by Gender." *American Sociological Review* 48:316–30.

Brown, Willie. 1999. "Breaking the Status Quo: San Francisco's Attempt to Ensure Fairness in Public Contracting." *Stanford Law and Policy Review* 10:227.

Bryk, Anthony S., and Stephen W. Raudenbush. 1992. *Hierarchical Linear Models: Applications and Data Analysis Methods*. Sage Publications.

Budig, Michelle J. 2002. "Male Advantage and the Gender Composition of Jobs: Who Rides the Glass Escalator?" *Social Problems* 49:258–77.

Cassirer, Naomi. 1996. "Race Composition and Earnings: Effects by Race, Region, and Gender." *Social Science Research* 25:375–99.

Chafetz, Janet S. 1984. *Sex and Advantage: A Comparative, Macro-Structural Theory of Sex Stratification*. Rowman & Allanheld.

Cohen, Philip N. 1998. "Black Concentration Effects on Black-White and Gender Inequality: Multilevel Analysis for U.S. Metropolitan Areas." *Social Forces* 77:207–29.

———. 2001. "Race, Class, and Labor Markets: The White Working Class and Racial Composition of U.S. Metropolitan Areas." *Social Science Research* 30:146–69.

Cohen, Philip N., and Matt L. Huffman. 2003. "Individuals, Jobs, and Labor Markets: The Devaluation of Women's Work." Department of Sociology, University of California at Irvine. Unpublished manuscript.

Cotter, David A., JoAnn DeFiore, Joan M. Hermsen, Brenda M. Kowalewski, and Reeve Vanneman. 1997. "All Women Benefit: The Macro-Level Effect of Occupational Integration on Gender Earnings Inequality." *American Sociological Review* 62:714–34.

———. 1998. "The Demand for Female Labor." *American Journal of Sociology* 103:1673–1712.

Cotter, David A., Joan M. Hermsen, and Reeve Vanneman. 1999. "Systems of Gender, Race, and Class Inequality: Multilevel Analyses." *Social Forces* 78:433–60.

Daspit, Nancy R. 1994. "The Family and Medical Leave Act of 1993: A Great Idea but a 'Rube Goldberg' Solution?" *Emory Law Journal* 43:1351.

Deaux, Kay. 1985. "Sex and Gender." *Annual Review of Psychology* 36:49–81.

Duncan, Otis D., and Beverly Duncan. 1955. "A Methodological Analysis of Segregation Indexes." *American Sociological Review* 20:210–17.

Edelman, Lauren B. 1990. "Legal Environments and Organizational Governance: The Expansion of Due Process in the American Workplace." *American Journal of Sociology* 95:1401–40.

England, Paula. 1992. *Comparable Worth: Theories and Evidence*. Aldine de Gruyter.

England, Paula, and Dana Dunn. 1988. "Evaluating Work and Comparable Worth." *Annual Review of Sociology* 14:227–48.

England, Paula, George Farkas, Barbara S. Kilbourne, and Thomas Dou. 1988. "Explaining Occupational Sex Segregation and Wages: Findings from a Model with Fixed Effects." *American Sociological Review* 53:544–58.

England, Paula, Melissa S. Herbert, Barbara S. Kilbourne, Lori L. Reid, and Lori M. Megdal. 1994. "The Gendered Valuation of Occupations and Skills: Earnings in 1980 Census Occupations." *Social Forces* 73:65–100.

England, Paula, Joan M. Hermsen, and David A. Cotter. 2000. "The Devaluation of Women's Work: A Comment on Tam." *American Journal of Sociology* 105:1741–51.

England, Paula, and Barbara S. Kilbourne. 1988. *Occupational Measures from the Dictionary of Occupational Titles for 1980 Census Detailed Occupations*. Interuniversity Consortium for Political and Social Research.

England, Paula, Lori L. Reid, and Barbara S. Kilbourne. 1996. "The Effect of the Sex Composition of Jobs on Starting Wages in an Organization: Findings from the NLSY." *Demography* 33:511–21.

Fossett, Mark A. 1988. "Community-Level Analyses of Racial Socioeconomic Inequality: A Cautionary Note." *Sociological Methods and Research* 16:454–91.

Gold, Michael E. 1993. *An Introduction to the Law of Employment Discrimination*. ILR Press.

Goldin, Claudia. 1990. *Understanding the Gender Gap: An Economic History of American Women*. Oxford University Press.

Grant, Don S., and Toby Parcel. 1990. "Revisiting Metropolitan Racial Inequality: The Case for the Resource Approach." *Social Forces* 68:1121–42.

Groshen, Erica L. 1991. "The Structure of the Female-Male Wage Differential: Is It Who You Are, What You Do, or Where You Work?" *Journal of Human Resources* 26:457–72.

Guo, Guang, and Hongxin Zhao. 2000. "Multilevel Modeling for Binary Data." *Annual Review of Sociology* 26:441–62.

Guthrie, Doug, and Louise M. Roth. 1999a. "The State, Courts, and Maternity Policies in U.S. Organizations: Specifying Institutional Mechanisms." *American Sociological Review* 64:41–63.

———. 1999b. "The State, Courts, and Equal Opportunities for Female CEOs in U.S. Organizations: Specifying Institutional Mechanisms." *Social Forces* 78:511–42.

Gutman, Arthur. 1993. *EEO Law and Personnel Practices*. Sage Publications.

Hannan, Michael T., and John H. Freeman. 1984. "Structural Inertia and Organizational Change." *American Sociological Review* 49:149–64.

Huffman, Matt L., and Steven C. Velasco. 1997. "When More Is Less: Sex Composition, Organizations, and Earnings in U.S. Firms." *Work and Occupations* 24:214–44.

Huffman, Matt L., Steven C. Velasco, and William T. Bielby. 1996. "Where Sex Composition Matters Most: Comparing the Effect of Job versus Occupational Sex Composition on Earnings." *Sociological Focus* 29:189–207.

Hultin, Mia, and Ryszard Szulkin. 1999. "Wages and Unequal Access to Organizational Power: An Empirical Test of Gender Discrimination." *Administrative Science Quarterly* 44:453–72.

Jacobs, Jerry A., and Mary Blair-Loy. 1996. "Race, Gender, Local Labor Markets and Occupational Devaluation." *Sociological Focus* 29:209–30.

Jacobs, Jerry A., and Ronnie J. Steinberg 1995. "Further Evidence on Compensating Differentials and Gender Gap in Wages." Pp. 93–124 in *Gender Inequality at Work*, edited by Jerry A. Jacobs. Sage Publications.

Jacobsen, Joyce P. 1994. *The Economics of Gender*. Blackwell.

Johnson, George, and Gary Solon. 1986. "Estimates of the Direct Effects of Comparable Worth Policy." *American Economic Review* 76:1117–25.

Kalleberg, Arne L., David Knoke, Peter V. Marsden, and Joe L. Spaeth. 1994. "The National Organizations Study: An Introduction and Overview." *American Behavioral Scientist* 37:860–71.

Kelly, Erin, and Frank Dobbin. 1999. "Civil Rights Law at Work: Sex Discrimination and the Rise of Maternity Leave Policies." *American Journal of Sociology* 105:455–92.

Kilbourne, Barbara S., Paula England, George Farkas, Kurt Beron, and Dorothea Weir. 1994. "Returns to Skill, Compensating Differentials, and Gender Bias: Effects of Occupational Characteristics on the Wages of White Women and Men." *American Journal of Sociology* 100:689–719.

Macpherson, David A., and Barry T. Hirsch. 1995. "Wages and Gender Composition: Why Do Women's Jobs Pay Less?" *Journal of Labor Economics* 13:426–71.

Major, Brenda, Dean McFarlin, and Diana Gagnon. 1984. "Overworked and Underpaid: On the Nature of Gender Differences in Personal Entitlement." *Journal of Personality and Social Psychology* 47:1399–1412.

Marsden, Peter V., Cynthia Cook, and David Knoke. 1994. "Measuring Organizational Structures and Environments." *American Behavioral Scientist* 37:891–910.

McArthur, Leslie. 1985. "Social Judgment Biases in Comparable Worth Analysis." Pp. 53–70 in *Comparable Worth: New Directions for Research*, edited by Heidi I. Hartmann. National Academy Press.

Milkman, Ruth. 1987. *Gender at Work: The Dynamics of Gender Segregation during World War II*. University of Illinois Press.

Nelson, Robert L., and William P. Bridges. 1999. *Legalizing Gender Inequality: Courts, Markets, and Unequal Pay for Women in America*. Cambridge University Press.

Parcel, Toby. 1989. "Comparable Worth, Occupational Labor Markets, and Occupational Earnings: Results from the 1980 Census." Pp. 134–52 in *Pay Equity: Empirical Inquiries*, edited by Robert T. Michael, Heidi I. Hartmann, and Brigid O'Farrell. National Academy Press.

Petersen, Trond, and Laurie A. Morgan. 1995. "Separate and Unequal: Occupation-Establishment Sex Segregation and the Gender Wage Gap." *American Journal of Sociology* 101:329–65.

Reid, Lori L. 1998. "Devaluing Women and Minorities: The Effects of Race/Ethnic and Sex Composition of Occupations on Wage Levels." *Work and Occupations* 25:511–36.

Reskin, Barbara F. 1993. "Sex Segregation in the Workplace." *Annual Review of Sociology* 19:241–70.

Reskin, Barbara F., and Debra B. McBrier. 2000. "Why Not Ascription? Organizations' Employment of Male and Female Managers." *American Sociological Review* 65:210–33.

Reskin, Barbara F., Debra B. McBrier, and Julie A. Kmec. 1999. "The Determinants and Consequences of Workplace Sex and Race Composition." *Annual Review of Sociology* 25:335–61.

Snijders, Tom, and Roel J. Bosker. 1994. "Modeled Variance in Two-Level Models." *Sociological Methods and Research* 22:342–63.

———. 1999. *Multilevel Analysis: An Introduction to Basic and Advanced Multilevel Modeling*. Sage Publications.

Spaeth, Joe L., and O'Rourke, Diane P. 1996. "Design of the National Organizations Study." Pp. 23–44 in *Organizations in America: Analyzing Their Structures and Human Resource Practices*, edited by Arne L. Kalleberg, David Knoke, Peter V. Marsden, and Joe L. Spaeth. Sage Publications.

Tam, Tony. 1997. "Sex Segregation and Occupational Gender Inequality in the United States: Devaluation or Specialized Training?" *American Journal of Sociology* 102:1652–92.

Tienda, Marta, and Ding-Tzann Lii. 1987. "Minority Concentration and Earnings Inequality: Blacks, Hispanics, and Asians Compared." *American Journal of Sociology* 93:141–65.

Tilly, Charles. 1998. *Durable Inequality*. University of California Press.

Tomaskovic-Devey, Donald. 1993a. "The Gender and Race Composition of Jobs and Male/Female, White/Black Pay Gaps." *Social Forces* 72:45–76.

———. 1993b. *Gender and Racial Inequality at Work: The Sources and Consequences of Job Segregation*. ILR Press.

———. 1995. "Sex Composition and Gendered Earnings Inequality: A Comparison of Job and Occupational Models." Pp. 23–56 in *Gender Inequality at Work*, edited by Jerry A. Jacobs. Sage Publications.

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

West, Candace, and Don H. Zimmerman. 1987. "Doing Gender." *Gender and Society* 1:125-51.