Working for the Woman? Female Managers and the Gender Wage Gap

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Most previous research on gender inequality and management has been concerned with the question of access to managerial jobs and the "glass ceiling." We offer the first large-scale analysis that turns this question around, asking whether the gender characteristics of managers—specifically, the gender composition and relative status of female managers—affect inequality for the nonmanagerial workers beneath them. Results from three-level hierarchical linear models, estimated on a unique nested data set drawn from the 2000 Census, suggest that greater representation of women in management does narrow the gender wage gap. Model predictions show, however, that the presence of high-status female managers has a much larger impact on gender wage inequality. We conclude that the promotion of women into management positions may benefit all women, but only if female managers reach relatively high-status positions.

The past two decades have seen substantial increases in the proportion of women in management. During this time, women's representation in managerial occupations increased from about one-third to one-half.[1] These positions confer well-documented benefits, including improved status, wages, autonomy, and overall work experience (England et al. 1994; Reskin and Ross 1992). In recent years, a spate of empirical research has addressed women's access to managerial authority (Blum, Fields, and Goodman 1994; Huffman and Cohen 2004a; Reskin and McBrier 2000; Smith 2002) and the "glass ceiling"—an unseen barrier between women and management or high-status positions (Cotter et al. 2001; Hultin 2003; Wright and Baxter 2000).

Although the question of access to managerial positions is critical to understanding persistent gender inequality in the labor market, the increase in women's managerial presence raises a broader question that is provocative and inherently sociological: What happens to the status of a subordinate group when some of its members attain positions from which they might reduce inequality? We use gender to gain insight into this question. Specifically, we ask whether the increase in women's representation in management "lifts all boats" by reducing gender inequality among *non*managerial workers or whether the benefits that accrue to female managers are limited only to those women. Clearly, the actions of managers affect those below them (Wright 1997). Yet, managers' role in reproducing gender inequality is conspicuously understudied, despite its relevance for persistent labor market inequality (Hultin and Szulkin 1999; Hultin and Szulkin 2003) and for broad-

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[1] This result is from our unpublished analysis of data from the Current Population Survey (available upon request). Interestingly, some of this change occurred at the same time that progress for U.S. women stalled on many other fronts (Cotter, Hermsen, and Vanneman 2004).

er questions about the status of subordinate groups.

We adopt a unique approach to this question. Rather than analyzing managerial workers alone (e.g., Huffman and Velasco 1997), or simply including them with nonmanagerial workers (e.g., Cohen and Huffman 2003a; Jacobs 1992), we analyze the wages of nonmanagerial workers as a function of the gender composition of their managers—the managers in their local industries. Further, we extend the existing literature by considering whether the *relative status* of female managers affects the pattern of gender inequality for the workers beneath them. If female managers influence gender inequality, the effect may depend on how highly placed those female managers are (Denmark 1993). Although this point may seem prosaic, careful attention to the relative status of female managers allows us to make more nuanced observations about the conditions under which labor market benefits extend to ascriptively similar subordinates. Using data from the 2000 U.S. Census, we estimate multilevel wage models with controls at three levels—the individual, the job, and the local industry—to assess the impact of female managers on gender inequality. These advances provide new leverage on theoretically important yet unanswered questions concerning the role of managers in gender stratification, with implications for the inequality trajectories of subordinate groups more generally.

MANAGERS AS AGENTS OF CHANGE OR COGS IN THE MACHINE?

Agents of Change

Many researchers believe that gender inequality at work results in part from the practices of managers—often assuming that these practices are associated with managers' gender. For example, Cotter and colleagues (1997:715) offer this as one reason why women benefit from occupational integration in the local labor market: "As more women in [positions of authority] make crucial decisions about salaries, promotions, hiring, and firing, gender differences in earnings *should* decline" (emphasis added). Similarly, Nelson and Bridges (1999) argue that the scarcity of women in authority positions sustains workplace gender inequality. Managers, by definition, have a certain level of organizational authority and thus might be poised to help reduce inequality at lower organizational levels. For female managers to reduce workplace inequality, however, two assumptions must hold. First, these women must be *motivated* to act in the interests of subordinate women. Second, they must have the *power* to influence outcomes for subordinates to affect gender inequality.

Gender creates a potential common interest between female managers and subordinates based on homophily (Ibarra 1992; McPherson, Smith-Lovin, and Cook 2001) or, in Kanter's (1977) term for the tendency of women to hire other women, "homosocial reproduction" (see also Elliott and Smith 2004; Pfeffer 1983). Sex similarity between subordinates and supervisors increases performance ratings by supervisors (Roth 2004; Tsui and O'Reilly 1989), and women's evaluation of potential female job candidates is less subject to pregnancy-related bias (Halpert, Wilson, and Hickman 1993).

More broadly, women express stronger support than men do for employer practices aimed at overcoming gender inequality. The 1996 General Social Survey asked for the level of agreement with the statement: "Because of past discrimination, employers should make special efforts to hire and promote qualified women." Employed women were 1.19 times more likely than men to agree (59.5 versus 49.8 percent, $p < .001$, N = 1,373). Among managers, the difference was larger, with women 1.32 times more likely than men to agree (53.5 versus 40.4 percent, $p = .068$, N = 193). Not only are women more supportive of efforts toward workplace equality in principle, but manager bias against such efforts among women is also less, suggesting that the presence of female managers should (if they have the power) promote gender equality.[2]

[2] The difference in agreement between nonmanagers and managers was not significant for women (60.4 percent versus 53.5 percent agreeing, $p = .197$, N = 753) but substantial for men (51.5 percent versus 40.4 percent, $p = .047$, N = 620). Baunach (2002) analyzed the same data set and reports no significant gender difference, but she used a subsample of respondents, with insufficient power to identify the difference (N = 313).

Ely (1995) argues that the presence of women in the upper echelons of organizations reduces the persistence of sex as a salient category for all workers, thereby weakening some of the negative consequences associated with gender imbalance (e.g., performance pressures, stereotyped role encapsulation, and exclusion from work-related networks). She demonstrates that the demographic composition of those holding powerful positions in organizations can have substantial effects on *all* workers, not just those holding positions at or near the top of organizational hierarchies.

Supporting this view, some empirical studies have shown less inequality where women occupy positions of authority. Hultin and Szulkin (2003) offer the most direct test of the wage effect of female managers, using rare employer–employee linked data from Swedish private-sector work establishments. They find a negative relationship between the gender wage gap among nonmanagerial workers and the proportion of women in managerial positions. This relationship, which remains substantial in the presence of controls for individual attributes, establishment characteristics, and industry, is present for both white- and blue-collar employees. Importantly, Hultin and Szulkin (2003) distinguish between higher-level decision makers (managers) and lower-level decision makers (supervisors). They find that the effect of the sex composition of *supervisors* on wage inequality is stronger than that of managers.[3] Although their analysis shows that the level of authority is an important consideration, female managers at high levels in the hierarchy are not shown to have a stronger effect than those at lower levels.[4]

A number of other studies are also consistent with the contention that female managers reduce gender inequality, but most rely on narrow samples or settings. An exception is a study of three U.S. cities, which finds that promotions from supervisor to manager occur more frequently under conditions of ascriptive similarity (race/ethnicity and gender) with immediate supervisors (Elliott and Smith 2004). Narrower studies have shown, for example, that California state agencies with more female managers exhibited less gender segregation in the 1970s and 1980s (Baron, Mittman, and Newman 1991), and savings and loans with women in management are more likely to hire women into managerial roles (Cohen, Broschak, and Haveman 1998). Additionally, Carrington and Troske (1995) find a strong link between the gender of business owners and the gender composition of their employees.

A series of studies investigating higher education settings shows that female administrators (Kulis 1997) or a female president (Pfeffer, Davis-Blake, and Julius 1995) are associated with less gender segregation (see also Konrad and Pfeffer 1991). In the legal profession, law firms whose corporate clients have many women in leadership positions show a greater increase in female partners (Beckman and Phillips 2005), and female decision makers tend to fill more vacancies with women (Gorman 2005). Finally, prime-time television shows with female producers, executive producers, and writers have a higher percentage of female major characters (Glascock 2001; Lauzen and Dozier 1999).

These studies imply that there is less gender inequality under conditions of greater female representation (and higher status) in management.[5] This may result from several distinct mechanisms, including increased access to organizational resources and power, homophily preferences, support for equity efforts, and weaker sex-based biases against female workers.

[3] Due to data limitations, this portion of their analysis was only performed on the blue-collar subsample.

[4] Although we do not link workers to managers directly, as do Hultin and Szulkin (2003), our study is unique in that it includes controls at three levels—the individual, the job, and the local industry. As such, it accounts for variation in gender inequality across larger social contexts, which purely organization-based analyses do not (Cohen and Huffman 2003b).

[5] Our emphasis on inequality effects differentiates this study from research on gender differences in leadership styles, but the pursuit of that question in the management and psychology literature provides some evidence of a less authoritarian orientation among female leaders, which may benefit female subordinates (for reviews, see Dobbins and Platz 1986; van Engen and Willemsen 2004).

Cogs in the Machine

Although there are reasons to believe female managers might reduce inequality, the underlying motivation and power assumptions are debatable. The motivations of female managers may be affected by two potential sources of loyalty or identity: their female peers in subordinate class positions and their managerial peers and superiors. Ely (1995) argues that to assume female managers are sympathetic to the women below them essentializes gender, while in practice gender is situationally enacted.[6] Class is one source of distinction that might prevent the expression of collective identity among women (Young 1994). In fact, a selection process may operate such that female workers are promoted into management partly for their affinity with the existing hierarchy. The disproportionate promotion of women who are "team players" may limit the potential for female managers to act against inequality.

Further, some women share men's biased views of women's work (Deaux 1985). For example, women and men in college similarly devalue the merit of female job applicants whose resumes reflect motherhood status (Correll and Benard 2005). With regard to networking, even if female or minority managers are likely to pass on job leads or other job-related information to subordinates, research suggests such contacts may not be systematically beneficial (Huffman and Torres 2002; Mouw 2003).

The managerial power assumption is also potentially flawed. It is not obvious that managers, especially those in bureaucratic organizations, are able to act autonomously on the basis of their own or women's interests. Instead they may be compelled to act under the mandates of routinization, efficiency, or profitability—or according to the prejudices of those higher up the hierarchy. This counterargument was summarized by Merton (1940:560) as, "authority . . . adheres in the office and not in the particular person who performs the official role." In fact, as Charles and Grusky (2004) show, the increase in managerial integration after the 1970s occurred during a period of growing bureaucratization, which implies limits on the discretionary power of lower-level managers. Kanter (1977) argues that female managers in particular occupy weak structural positions. In Ridgeway's words, they are "handicapped by their lower power and by interactional gender mechanisms" (1997:227). Affirmative action programs have been more successful at integrating lower and middle levels of management (Brenner, Tomkiewicz, and Schein 1989), and to some extent women's increasing managerial presence reflects "title inflation" (see Jacobs 1992)—the reclassification of previously nonmanagerial workers as managers with little increase in pay or authority.

Finally, it is possible that due to a baseline sectoral segregation, women are typically managers in workplaces with lower quality jobs. Both Shenhav and Haberfeld (1992) and Pfeffer and Davis-Blake (1987) find lower earnings for both men and women in workplaces with more female managers or administrators. If female managers are concentrated in organizations with more female workers, then any positive effect of female manager attitudes or behavior may be swamped by negative gender concentration effects.

In summary, female managers may enhance the labor market prospects of the women who work below them. Their homophilous preferences or affiliations might promote equality, and they may have less to gain from discrimination and therefore be more motivated to help other women. Additionally, women may be more aware than men of discriminatory practices and less susceptible to cognitive processes leading to gender bias. Any of these processes may smooth the social and organizational path of female subordinates. In contrast, bureaucracy, market pressures, divided loyalties, past discrimination, or the mandates of those more powerful may render the ascriptive characteristics of managers largely moot with regard to inequality.

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[6] Research in psychology shows that to the extent there are gender differences in moral reasoning, as advanced by Gilligan (1982), they are context-dependent (Ryan, David, and Reynolds 2004) and conditioned on, among other factors, socioeconomic status. For our purposes, however, we note that Jaffee and Hyde (2000) find greater gender differences in moral reasoning at higher levels of social class.

Managers' Relative Status

Regardless of how female managers might reduce inequality, their impact may depend critically on their status (Denmark 1993). Even highly motivated female managers working together may not be able to influence gender inequality if they are relatively powerless. Both the identity/loyalty issue and the question of managerial power highlight the possibility of substantial interactions between the representation of women and their relative status. Although disparate studies have investigated the effect of female representation among managers at different levels (e.g., university presidents, screenwriters, law firm partners), Hultin and Szulkin's (2003) is the only one to directly test for the effect across different levels within one setting, and their ability to address the issue is limited by their data, which includes only undifferentiated manager and supervisor categories. In contrast, we use a continuous measure of vertical segregation (explained below) to tap the relative status of female managers.

Local Industries

Most of the research in this area is concerned with the direct effects of managers within organizations. Studies have thus been designed to draw from managers and workers who are as closely linked as possible, exemplified by the work of Hultin and Szulkin (2003) and Elliott and Smith (2004). Women in positions of authority, however, may change gender dynamics at various levels of proximity: among immediate subordinates, within their organizations in general, across organizations, and across larger social contexts. We know that gender inequality varies systematically across larger social contexts, including metropolitan labor markets (Cohen and Huffman 2003a; Cotter et al. 1997) and national industries (Fields and Wolff 1995; Wharton 1986). This variation is not captured when analysis is limited to direct examination of organizations (Cohen and Huffman 2003b).

One way the processes reproducing inequality across contextual levels would be linked is if female managers in one workplace affect proximate organizations. For example, an employer's decision to hire women creates a competitive advantage relative to those who do not (Ashenfelter and Hannan 1986), especially when women are paid less than men (Neumark and Stock 2006). Such competition is one way that corporate practices—which presumably include those related to gender inequality—are adapted by different actors within an organizational field. Additionally, employment practices can be adopted by organizations in an effort to increase legitimacy or to appear in compliance with a changing legal environment (DiMaggio and Powell 1983; Tomaskovic-Devey and Stainback 2007).

DiMaggio and Powell's (1983) theory has prompted an extensive literature on how to operationalize the reference groups or fields for organizational behavior (Greve 2005; Massini, Lewin, and Greve 2005; Strang and Soule 1998). Employing organizations are part of labor markets that are local (represented by metropolitan areas), industries that are national (represented by categorization schemes with various degrees of detail), and the intersection of the two: *local industries.* We believe the local industry—the aggregate of organizations that produce a common product within a common local labor market—is an appropriate starting point for our questions. This approach draws from work on local industrial dominance by South and Xu (1990), who argue that local industries are sites of intersection between structural forces and individual attainment processes. These units combine functional commonality with social proximity, capturing the interaction of these fields.

We thus expect local industries to display less internal variation in gender-related practices than either metropolitan labor markets or national industries as a whole. This implies that a given restaurant, for example, is likely to resemble other restaurants in its local area more than it resembles either restaurants in the entire country or all employers in the local labor market. Although we cannot fully demonstrate this pattern, we provide a simple illustration based on one key variable for establishments: the gender composition of managers.[7]

Using data from all large U.S. private-sector firms collected by the U.S. Equal Employment Opportunity Commission (EEOC) in 2002, we

[7] This is the measure used by Ashenfelter and Hannan (1986), who examine the pattern of gender representation in management for banks across local markets.

measured the similarity of establishments on managerial gender composition across 275 metropolitan labor markets, 301 national industries, and 10,131 local-industry cells formed by the intersection of labor markets and industries. First, we calculated the natural logarithm of the percentage female among "officers and managers" at each establishment. We then calculated the standard deviation in the natural logarithm for each contextual unit—each metropolitan area, national industry, and local industry. Finally, we averaged the standard deviations across these contextual units. The means of these standard deviations are 1.45 for metropolitan areas, 1.22 for national industries, and 1.03 for local industries.[8] In terms of the gender composition of managers, establishments on average are indeed more similar to others within their local industries than they are to those within their entire local labor markets or national industries as a whole. This supports the presumption that gender-related organizational dynamics generally cluster or reproduce more tightly within local industries than within these larger units.

Of course, we remain interested in within-organization effects of managerial gender composition on gender inequality. Despite reasons to suspect larger processes, this more direct effect remains the most plausible pathway by which female managers influence gendered outcomes, as has been shown in the limited research that uses such linked data. In the absence of such data on a generalizable scale, however, we take heart from the results of our EEOC exercise, which imply that for a given worker, the gender composition of the local industry's managers is likely to approximate that of her own establishment, at least compared to the correspondence obtained at the level of the local labor market or national industry. We thus consider local-industry managerial composition as a proxy for establishment management characteristics.

HYPOTHESES

Previous research has shown that the overall wage gap largely results from between-job and between-occupation inequality, whereby female-dominated jobs and occupations pay less than otherwise comparable male-dominated lines of work (Cohen and Huffman 2003a; England et al. 1994; Huffman and Velasco 1997; Tam 1997), in part because female-dominated jobs offer fewer training opportunities (Tomaskovic-Devey and Skaggs 2002). However, inequality within job and occupational categories also contributes to wage inequality. Clearly, managerial composition and relative status could shape wage inequality through either route, as managers might influence both sorting and the training and rewards processes. Rather than make predictions without justification from prior theory or research, our models account for both possibilities. We also do not consider more complex interactions (e.g., McCall 2001), such as those involving the conditional effects of managers' race and gender.

Our analysis concerns the extent to which wages for men and women are sensitive to both the *representation* of women among managers and the *relative status* of those female managers. We test two hypotheses, beginning with whether the gender wage gap is smaller in local industries where there are more female managers. Specifically, we test:

Hypothesis 1: There is less gender wage inequality in local industries with a higher proportion of female managers.

This hypothesis addresses the association between the *representation* of women in managerial positions and wage inequality. If female managers tend to cluster at the bottom of managerial hierarchies, though, their mere representation in management may be insufficient to alter wage inequality. Therefore, we test the interaction between the representation and the relative status of female managers:

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[8] Our data are from the 2003 EEOC files, based on required filings by all private-sector establishments with 50 or more employees and smaller firms if they are federal contractors. The calculations are based on about 170,000 individual establishments with any managerial workers located in identifiable metropolitan areas (as used below). We set logged percent female to 0 where there were less than 1 percent female managers (the results were substantively the same when we used unlogged percentages). The differences in mean standard deviations were highly significant at conventional levels. Details are available from the authors. See Cohen and Huffman (2007) and Robinson and colleagues (2005) for more recent analyses of this data set.

Hypothesis 2: There is less gender wage inequality in local industries (a) that have more female managers and (b) in which female managers on average hold higher-status positions.

To test these hypotheses, one must combine data from several sources and organize them into a multilevel structure. Despite large data sets and sophisticated methods, as is typical in large-scale studies of labor market inequality, we cannot offer causal tests of these hypotheses. Rather, we test whether the data are consistent with the patterns predicted by these hypotheses. Following Reskin (2003:14), we believe that "although contextual effects are not themselves mechanisms, they are proxies for mechanisms that vary across settings." Next, we describe our data collection and manipulation as well as our statistical modeling strategy.

DATA, MEASURES, AND MODELS

DATA

We investigate our hypotheses by analyzing data at three conceptual levels: individual workers, jobs, and local industries. We nest individual workers in "jobs," defined as three-digit occupation by three-digit industry by metropolitan area cells (Cohen and Huffman 2003a; Huffman and Cohen 2004b). Our innovation is that employed respondents are separated into managerial and nonmanagerial jobs. Nonmanagerial workers are the subjects of our wage analysis. Each nonmanagerial job is nested within a local industry—a three-digit industry in a metropolitan labor market. Manager characteristics are drawn from the managerial workers in each local industry and used as independent variables.

Our primary data source is the combined 2000 Census 5- and 1-percent Public Use Microdata Samples (PUMS). We analyze the wages of metropolitan nonmanagerial civilian workers ages 25 to 54 years. We restrict our sample to those who are employed in jobs with at least 10 people and local industries with at least 10 managers in each of at least two managerial occupations—this is necessary for calculating female managers' relative status (see below). This process yields a sample of approximately 1.32 million workers nested in 29,294 local jobs, which are in turn nested in 1,318 local industries (representing 155 industries in 79 metropolitan labor markets).[9] Additional measures for characteristics of local industries, based on their metropolitan areas, are drawn from the Census 2000 Summary Files. The *Dictionary of Occupational Titles* (DOT), which includes measures of occupational skills and requirements, is the final source of data (Tomaskovic-Devey and Skaggs 2002). Appending these measures to occupations in the 2000 Census required converting occupation codes from the 1990 scheme to the new coding scheme employed in 2000.[10]

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9 We exclude workers who were self-employed, in military-specific occupations, or in the armed forces because their wages and promotions are not determined by local managers. We also exclude those with wages outside the range of $1 to $300 per hour; legislators, for whom we have no occupational characteristics; and those with no specific metropolitan area identified (mostly rural workers). The age restriction is applied after job cell characteristics are calculated (so that all workers contribute to job characteristics, not only those for whom the outcome is analyzed). The job and local industry restrictions reduce the final sample from 3.45 to 1.32 million workers, mostly by removing workers from smaller labor markets. Notable differences in the final sample include more foreign-born workers, more concentration in the Northeast and West regions, and more highly educated workers. Workers excluded by job and local industry restrictions have logged wages .18 lower than those in the final sample, but this is reduced to less than .05 when adjusted for observed individual characteristics. We have no reason to suspect that our selection criteria introduce systematic biases with regard to our hypotheses, but we cannot rule out that possibility.

10 The DOT database contains information for nearly 13,000 occupations corresponding to about 500 occupations in the three-digit codes used by the 1990 Census. We matched 2000 Census occupations to DOT occupations using a crosswalk file from the National Occupational Information Coordinating Committee and calculated mean scores across DOT occupations for each Census occupation. We used NOICC Master Crosswalk v. 4.3 (revised November 15, 1999), and a file titled "DOTCEN00" ("Crosswalk linking the 2000 Census occupations to those from the Dictionary of Occupational Titles"), both accessed from the National Crosswalk Service Center (http://www.xwalkcenter.org) on March 10, 2006.

MEASURES

LEVEL 1: INDIVIDUALS. In our models, the dependent variable is the natural logarithm of the *hourly wage,* which is annual earnings divided by hours worked. At the individual level, we include a binary variable for *gender* (female = 1) and four dummy variables to represent respondents' *ethnicity* (coded 1 if the respondent is Black, Latino, Asian, or other ethnicity).[11] White is the omitted race/ethnicity. We use dummy variables to capture differences in *educational attainment* (less than high school, some college, bachelor's degree, or master's degree or above), whether the respondent is *married, formerly married* (divorced, widowed, or separated), and *foreign born.* We also include dummy variables to control for *disability status* and the presence of *children younger than six years* in the household, as well as whether the respondents *do not speak English well* and whether they are *currently attending school.* Continuous variables measure *potential labor market experience* (age minus years of education minus 5) and its *square,* in addition to *number of own children* in the household.

LEVEL 2: JOBS. At the job level, our independent variable of primary interest is *percent female.* To account for nonlinear effects of percent female, we include *percent female squared.*[12] We measure other demographic characteristics of jobs, including *percent Black, percent Latino,* and *percent Asian.* We also control for the *percent part-time* employed in the job. From the DOT, we include three measures: *standard vocational preparation* (SVP), *general educational development* (GED), and *physical strength* (STR). SVP measures the amount of training time needed to learn the techniques and obtain the information necessary for average job performance (high values of this scale represent a longer period of time required to acquire the skills). SVP can be thought of as a measure of occupation-specific human capital (Tomaskovic-Devey and Skaggs 2002), while GED measures "the typical requirement of the occupation for schooling that is not vocationally specific" (England, Hermsen, and Cotter 2000:1742). GED is calculated as the mean of values required for mathematics, language, and reasoning preparation. Finally, STR is coded to reflect strength requirements for each occupation, ranging from 1 (sedentary) to 5 (very heavy). This variable is intended to capture the manual nature of occupations, which features prominently in the gender division of labor (Charles and Grusky 2004).

LEVEL 3: LOCAL INDUSTRIES. At the local industry level, our key independent variables are *percent female* and the *relative status* of female managers. In the absence of a direct measure of decision-making authority, we identify managers as those in "management occupations" in the occupational classifications of the federal government (U.S. Census Bureau 2003). This classification clearly is a relevant, if imperfect, measure of authority.[13] Percent female among

[11] Because of overlapping racial and ethnic identification in the 2000 Census, we use a descending selection to reach mutually exclusive categories, in the following order: Latino, Black, Asian, other, White. Latinos are thus coded as such regardless of their responses to the race question, and Whites are those who selected no Latino ethnicity or other race. This conforms to the recommendation of the federal Office of Management and Budget with regard to civil rights enforcement (Goldstein and Morning 2002).

[12] Consistent with many findings in the literature (e.g., Cohen and Huffman 2003a; Cotter et al. 1998; England et al. 1994), we found a linear effect on wages of the gender composition of jobs in models with controls at all levels. Cotter and colleagues (2004), however, present evidence from the 2000 Census that calls this simple relationship into question: cubic and 4th-power fits for women and men, respectively, in the bivariate relationship between occupation percent female and median earnings at the national level. Closer examination revealed a 4th-power fit in the bivariate relationship between job gender composition and average wages. With the introduction of controls at the individual level (which also captures women's lower average individual wages), the best fit was cubic, and with controls at the job level added, the best fit fell to quadratic. This did not change with the addition of local-industry controls. Therefore, we model the job gender composition effect as quadratic.

[13] We checked the Multi-City Study of Urban Inequality for three large U.S. cities and found that 65 percent of workers in managerial occupations report having the authority to hire and fire others

managers in each local industry is simply obtained from the PUMS files, using the same criteria for selection we use for nonmanagerial workers at level 1.

To measure female managers' status relative to male managers—our proxy for decision-making power—we use a measure of vertical segregation, the Index of Net Difference (ND), as described by Lieberson (1976). ND is the difference between the likelihood that a randomly chosen man is employed in a higher-status managerial occupation than a randomly chosen woman and the opposing probability, that a randomly selected woman works in a higher-status managerial occupation than a randomly chosen man. Specifically, ND is given by:

$$\text{ND} = 100 \times (\Sigma \text{M}_i \text{CF}_i - \Sigma \text{F}_i \text{CM}_i)$$

M_i and F_i equal the proportion of males and females, respectively, in managerial occupation *i*. CF_i equals the cumulative proportion of females in managerial occupations ranked below managerial occupation *i,* and CM_i equals the analogous cumulative proportion of men. When ND equals zero, men and women are equally likely to occupy high-status occupations. ND equals 1 when all women are in higher-status occupations than men, and when ND equals –1 all women are in lower-status positions than men.

Calculating ND requires an ordinal ranking of managerial occupations. Unfortunately, authority over other workers, although theoretically central to analyses of class dynamics at work (Wright 1997), is not a prominent concern in most occupational research. We know of no source that reports the relative decision-making authority or status of managerial occupations. Despite detailed descriptions of thousands of occupations, neither the DOT nor the more recent federal O*Net descriptor system measure authority directly. One approach is to dichotomously code occupations from the Census as having authority or not, based on the appearance of the words "manager," "supervisor," or "administration" in the occupation title (England et al. 1994). This does not, however, rank occupations according to the level of authority—chief executives and food service managers, for example, are both simply coded as possessing authority.

In the absence of a direct measure of authority, we first select occupations in the federal system's "management occupations" category, then apply a measure based on both skills and training (representing expertise) and earnings (representing recognition and rewards). Our measure of the status of each managerial occupation is the average of two factors: (1) the average of z-scores for years of education, SVP, and GED and (2) the z-score for earnings. Managerial occupations are thus ranked according to equal weightings of the skills and training required (education, SVP, and GED) and also average earnings.[14] In the resulting authority ranking, natural-science managers are highest and gaming managers are lowest.[15]

We control for other important characteristics of local industries that could affect gender inequality. Among managers, we include *managers as a percentage of all workers.* This is intended to capture the level of rationalization or bureaucratization of the local industry, which are traits associated with reduced reliance on ascriptive characteristics in determining work positions and rewards (Jackson 1998; Reskin 2003).

We also include characteristics of local labor markets at this level, drawing data from the PUMS and 2000 Census Summary Files.[16] To

(compared to 30 percent of those in other occupations) and 48 percent reported influencing the rate of pay of others (compared to 22 percent of those in nonmanagerial occupations).

[14] Some managerial occupations are specific to one industry (e.g., funeral directors), but most occur across at least several industries (e.g., human resource managers and chief executives). To calculate status, we combined managers from all industries in all labor markets.

[15] Because women are heavily represented in some higher-status managerial occupations (e.g., medical and health service managers and educational administrators) and poorly represented in some lower-status positions (e.g., construction managers), our status measure and percent female are only weakly correlated ($r = -.16$). Across all local industries, female managers' relative status is higher where their representation in management is greater ($r = .09$).

[16] In principle, labor market variables could constitute a fourth level of the analysis. However, because our hypotheses do not focus on local labor market dynamics per se, we include these variables at level

capture other aspects of local gender dynamics we control for the overall level of *occupational gender segregation* and the *demand for female labor.* Both of these measures are computed across the 33 major occupational categories reported in the Summary Files. The segregation measure uses the index of dissimilarity (Duncan and Duncan 1955). This measure has a significant association with the gender gap in earnings across U.S. labor markets, such that those with higher levels of segregation have lower relative wages for women, regardless of whether they work in male- or female-dominated occupations or jobs (Cohen and Huffman 2003b; Cotter et al. 1997). The demand measure reflects the number of women who would be employed if local occupations had the gender composition observed nationally, divided by the actual number of employed women. Labor markets with higher levels of demand for female labor also have significantly less gender wage inequality (Cotter et al. 1998). Our measure is similar to that used by Cotter and colleagues (1998), although with less occupational detail. These two controls are especially important if we are to avoid spurious effects whereby wage gaps and women's managerial representation are both more favorable as a result of larger factors affecting all women in the local labor market.[17]

To capture other local economic conditions, we include the *unemployment rate* and industrial composition, measured by *percentage in manufacturing* among all workers. Region is controlled with dummy variables for the *South, West,* and *Midwest* (Northeast is omitted). We also include demographic characteristics: the natural logarithm of *population size,* the *percentage Black* and *percentage Latino,* and the rate of *in-migration,* measured by the percentage of local residents who moved to the metropolitan area in the previous five years.

Descriptive statistics at each level of the analysis appear in Table 1, which shows that men in our sample have average logged wages of 2.84 ($17.13), compared to 2.67 ($14.42) for women, for a gender wage ratio of .84. Managerial composition ranges from 2 to 90 percent, with means of 34 percent for men and 48 percent for women. Female managers' relative status has a mean of –.075 and ranges from –.72 to .79.

Illustrative Examples

Several examples clarify our data structure and its applicability to our hypotheses. Table 2 shows a comparison across four local industries: restaurants and computer systems in the Los Angeles and New York metropolitan areas.[18] Our sample has 1,887 managers in Los Angeles restaurants. Of these, food service managers are the most common (N = 1,450). This local industry also has 44 nonmanagerial occupations (jobs) with 10 workers or more and a total of 10,422 workers, the plurality of whom are cooks (N = 2,928).

Our analytic strategy works on the assumption that the behavior of managers in the local industry—which is influenced by their gender composition and the relative status of the women among them—may influence the wages of the nonmanagerial workers under them. In the restaurant example, we see that there are more female managers in Los Angeles (42.4 percent) than in New York (33.4 percent). The female managers in New York, however, have higher relative status than those in Los Angeles, mostly because they are less concentrated in the low-status food service manager occupation. The ND score for New York (.021) is thus higher than that for Los Angeles (–.047). Among nonmanagerial workers, New York has fewer women (40.7 versus 45.4 percent), and the gender wage ratio among nonmanagerial workers is higher in New York, where women earn 95 percent of

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3, which allows us to use the HLM software (see below) for computing and greatly simplifies computation and interpretation.

[17] In our sample, female representation in management is negatively correlated with occupational segregation ($r = -.06$, $p < .05$) but not with demand. Female managers' relative status, on the other hand, is higher in local labor markets with more occupational segregation ($r = .06$, $p < .05$) and lower in those with more female demand ($r = -.08$, $p < .01$). This implies that in more integrated markets, and those with more female-dominated occupational structures, women are more likely both to be managers and to be crowded into low-level managerial positions.

[18] We chose the restaurant industry because it has the most local cases in our data set and the computer system industry as a male-dominated contrast.

Table 1. Descriptive Statistics for Variables Used in the Analysis (by gender)

	Men	Women	Min.	Max.
Individuals				
Hourly Wage (natural log)	2.84	2.67	0	5.70
Less than High School	.18	.11	0	1
High School Complete	.23	.22	0	1
Some College	.27	.32	0	1
B.A.	.20	.23	0	1
M.A. or Higher	.12	.13	0	1
In School	.08	.10	0	1
Work Disability	.17	.15	0	1
White	.58	.60	0	1
Black	.11	.15	0	1
Asian	.08	.08	0	1
Latino	.20	.15	0	1
Other Race/Ethnicity	.02	.02	0	1
Never Married	.29	.25	0	1
Married	.57	.56	0	1
Formerly Married	.14	.18	0	1
Own Children in Household	.90	.90	0	12
Children Under 6 in Household	.23	.19	0	1
English (not spoken well)	.17	.11	0	1
Foreign Born	.30	.24	0	1
Potential Experience	18.89	19.23	0	48
Potential Experience Squared	432.73	448.85	0	2304
Jobs				
Proportion Female	.28	.69	0	1
Proportion Part-Time	.13	.23	0	1
Proportion Black	.11	.15	0	1
Proportion Latino	.20	.15	0	1
Proportion Asian	.07	.08	0	1
General Educational Development	3.26	3.54	1.00	6
Specific Vocational Preparation	5.66	5.64	1.00	8.22
Strength	2.27	1.82	1.00	5
Local Industries				
Manager Percent Female	33.85	47.80	2.33	90.00
Female Manager ND	–.07	–.08	–.72	.79
ND × Percent Female Interaction	–2.52	–3.43	–51.72	45.21
Percent Managers	10.19	9.95	1.20	37.72
MA Gender Segregation	.38	.38	.34	.48
MA Female Labor Demand/Supply	1.01	1.01	.92	1.04
Northeast	.26	.30	0	1
Midwest	.15	.14	0	1
South	.26	.27	0	1
West	.32	.29	0	1
MA Population (ln)	15.81	15.87	11.92	16.87
MA Percent Black	13.20	13.59	.49	39.95
MA Percent Latino	16.70	16.32	.63	47.66
MA Unemployment	5.89	5.89	3.45	11.98
MA Percent in Manufacturing	12.42	12.28	3.66	39.41
MA Percent In-migration	30.02	29.75	22.74	35.62

Source: 2000 U.S. Census Public Use Microdata files and other sources (see text).
Notes: MAs are (Consolidated) Metropolitan Statistical Areas. ND is the index of net difference, showing women's status relative to men; 1 = all men in top positions and 1 = all women in top positions (see text).
All gender differences are significant at $p < .001$, except Asian (n.s.) and own children in household ($p < .05$).

Table 2. Managerial and Nonmanagerial Occupations in Four Local Industries

	Los Angeles			New York			T-tests	
	N	*Percent Female*	Wage	*N*	*Percent Female*	Wage	*Percent Female*	Wage
Restaurants and Other Food Services								
Largest Managerial Occupations								
Chief Executives	31	22.6	64.30	30	30.0	65.06		
General & Operations Managers	179	30.2	19.30	154	29.9	19.75		
Marketing & Sales Managers	61	44.3	30.97	26	50.0	21.90		
Human Resources Managers	110	48.2	15.33	65	44.6	12.41		
Food Service Managers	1,450	43.7	15.27	1,681	32.9	18.46	*	*
Total (including occupations not shown)	1,887	42.4	17.33	2,014	33.4	19.31	*	*
Gender ND		–.047			.021			
Largest Nonmanagerial Occupations								
Cashiers	827	80.9	11.00	701	78.3	11.24		
Waiters & Waitresses	2,529	66.0	12.18	2,824	61.6	12.44	*	
Food Preparation Workers	444	48.0	10.18	551	42.7	10.17		
Supervisors, Food Prep., & Service Workers	717	49.9	14.41	627	40.8	14.01	*	
Bartenders	304	35.9	12.86	452	46.7	13.62	*	
Cooks	2,928	29.5	11.88	2,384	18.8	11.12	*	
Attendants & Bartender Helpers	401	14.5	10.12	303	27.4	9.82	*	
Chefs & Head Cooks	482	14.0	14.04	1,215	11.1	14.19		
Dishwashers	255	13.7	9.23	305	11.2	8.42		
Driver/Sales Workers & Truck Drivers	316	13.0	11.98	237	8.0	12.02		
Total (including occupations not shown)	10,422	45.4	12.19	11,025	40.7	12.24	*	
Women	4,730		11.66	4,487		11.87		
Men	5,692		12.65	6,538		12.50		
Gender Wage Gap			92.2			95.0		
Computer Systems Design and Related Services								
Largest Managerial Occupations								
Chief Executives	89	18.0	55.35	108	18.5	63.40		
General and Operations Managers	45	22.2	44.13	57	19.3	48.39		
Computer & Information Systems Managers	115	25.2	32.36	216	24.5	40.28		*
Financial Managers	24	45.8	32.18	44	38.6	38.19		
Marketing & Sales Managers	84	36.9	30.69	132	39.4	40.92		*
Total (including occupations not shown)	589	28.9	37.49	893	31.5	43.06		*
Gender ND		–.140			–.196			
Largest Nonmanagerial Occupations								
Sales Representatives	87	27.6	37.41	131	31.3	43.24		
Computer Support Specialists	92	31.5	24.12	128	25.8	27.21		
Computer Programmers	352	23.9	33.40	765	24.2	34.30		
Computer Software Engineers	428	19.6	34.03	665	22.4	37.92		*
Computer Scientists & Systems Analysts	296	20.6	31.10	617	20.1	36.10		*
Network Systems & Data Comm. Analysts	155	17.4	21.99	238	17.2	33.65		*
Total (including occupations not shown)	2,104	30.0	28.82	3,586	29.8	32.94		*
Women	632		25.33		1,069	28.47		*
Men	1,472		30.32		2,517	34.83		*
Gender Wage Gap			83.5			81.7		

Source: 2000 U.S. Census Public Use Microdata files.
Notes: Full names are Los Angeles-Riverside-Orange County and New York-Northern New Jersey-Long Island. Managerial occupations sorted by status; nonmanagerial occupations sorted by percent female. Wages are means. ND is the index of net difference (see text). * $p < .05$ (two-tailed).

men's average wages, compared to 92.2 percent in Los Angeles.

The computer systems industry in these two labor markets is much more male dominated, has higher wages than restaurants, and features more striking gender inequality. In this case, female managers are more common in New York than in Los Angeles, but New York has more women in lower-status marketing manager jobs and fewer in higher status positions. Among nonmanagerial workers, the gender wage ratio is lower in New York than in Los Angeles (.82 versus .84). The similar gender distribution across jobs suggests this gap mostly reflects inequality within jobs rather than job segregation.

Continuing the restaurant example, our data set is adequate to analyze 58 local industries, that is, the restaurant industry in 58 local labor markets. Figure 1 shows the relationship between manager percent female and the gender wage ratio among nonmanagerial workers in these local industries, with the largest metropolitan areas highlighted. For illustration, we split the local industries at the median gender ND score (–.047) and show each half in a separate plot. In the low-ND markets—where female managers are in lower-status positions relative to male managers—there is no relationship between percent female and the gender wage ratio. In the high-ND markets, however, local industries with more women in management exhibit a smaller gender gap. In this industry, the pattern across labor markets suggests that the effect of higher female representation among managers may be conditional on their attainment of higher-status managerial positions.

These examples also illustrate an issue we mentioned above—that female managers may be ghettoized in the same industries where female workers are concentrated, producing a negative relationship between female managerial concentration and average wages for both female and male workers.[19] Table 3 displays the distribution of male and female workers across categories of female manager representation, as well as median wages and female managers' relative status. The table shows that 31.8 percent of men, but only 6.5 percent of women, work in local industries with fewer than 20 percent female managers, typified by the construction industry. On the other hand, more than one-quarter of women, but only 8 percent of men, work in local industries where 60 percent or more of the managers are female, typified by hospitals. Thus, the gender of workers and managers is clearly related. Note that the most common situation involves female managerial representation between 20 and 40 percent (e.g., restaurants). This is the group of local industries in which female managers have the lowest relative status (ND = –.14) and the gender gap in median wages is greatest, with nonmanagerial women earning just 75 percent of the male wage. We now investigate these relationships on a larger scale, taking into account possible confounding factors at the levels of the person, job, and local industry.

Statistical Models

Our data have a nested structure, with individual workers nested within jobs, which in turn are nested within local industries. We therefore use three-level hierarchical linear models (Raudenbush and Bryk 2002), which are justified both on statistical and substantive grounds. First, hierarchical models avoid downwardly biased estimation of standard errors when data are nested (Guo and Zhao 2000; Raudenbush and Bryk 2002). Additionally, they provide the flexibility to specify cross-level interaction effects, on which our hypotheses are based. For example, is the individual gender effect stronger or weaker depending on the gender composition of the job and its local managers?

Specifically, in our individual-level model, logged wages are a function of a model intercept (average wages for men), a female effect (within-job gender inequality), and controls. Formally, our individual-level model can be expressed as:

$$Y_{ijk} = \pi_{0jk} + \pi_{1jk}(\text{female}_{ijk}) + \pi_{2jk}a_{1ijk} + \ldots + \pi_{mjk}a_{mijk} + e_{ijk}$$

where Y_{ijk} is the logged wage of person i in job j in local industry k, and π_{0jk} is the inter-

[19] In fact, the data show a small positive bivariate correlation between female managerial representation and (logged) wages for both men (r = .09) and women (r = .07). A closer inspection, though, shows this is because a cluster of male-dominated blue-collar jobs, such as construction, have relatively low wages and almost no female managers.

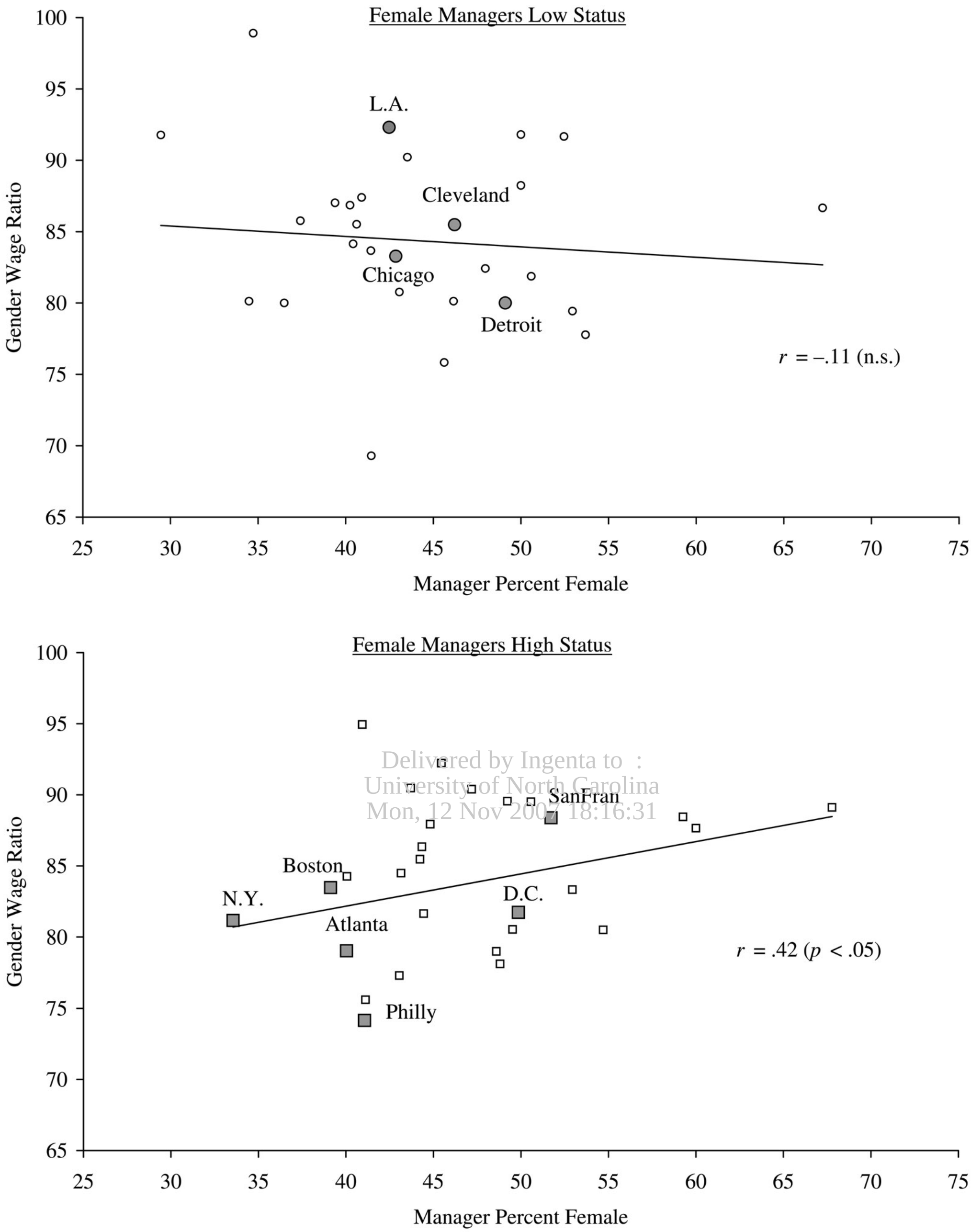

Figure 1. Restaurant Industry Manager Percent Female and the Gender Wage Gap Among Nonmanagerial Workers, by Female Managers' Relative Status in 58 Labor Markets

Notes: Local industries split at median ND (–.047), N = 29 in each figure. Correlations and regression lines weighted by sample size. N's range from 51 (Santa Barbara, CA) to 2,886 (New York); the largest 10 labor markets are labeled. Wage ratio is women's median wage as percent of men's.

cept for job j in local industry k. Next, π_{1jk} is the individual-level effect of gender, a_{mijk} denote the M individual-level control variables, and π_{2jk} through π_{mjk} are the associated individual-level regression coefficients. Finally, e_{ijk} is the level-1 random effect. The coefficient π_{1jk} is of particular interest because it represents the net within-job wage

Table 3. Nonmanagerial Worker Distribution and Wages, by Manager Percent Female in Local Industries

Manager Percent Female	Median ND	Worker Distribution Total	Men	Women	Median Wage Total	Men	Women	Wage Ratio	Most Common Industry
≤ 19	–.022	19.4	31.8	6.5	15.00	15.38	12.88	.84	Construction
20 to 39	–.140	25.7	31.0	20.3	16.92	19.23	14.42	.75	Restaurants
40 to 59	–.074	37.7	29.2	46.5	15.63	17.79	14.51	.82	K–12 Schools
≥ 60	–.050	17.2	8.0	26.8	15.38	17.00	14.96	.88	Hospitals
Total	–.068	100.0	100.0	100.0	15.77	17.31	14.42	.83	

Source: 2000 U.S. Census Public Use Microdata files.
Notes: Worker distribution and wages are measured at the individual level. Manager ND is the median for local industries within each category, weighted by the number of workers in each local industry. Most common industries are those with the most total workers in each category, shown for illustration. ND is the index of net difference (see text).

gap. To simplify interpretation of the results, all variables except gender are centered around their grand means. Thus, π_{0jk} represents the average wage for men at the mean of the control variables.

Our job-level model estimates both the level-1 intercept and the level-1 female effect as a function of job percent female (and its square) and our job-level controls. Each level-1 coefficient relating individual characteristics to wages can be modeled as either a fixed or random effect across jobs. We allow only the level-1 intercept and the coefficient for the female dummy variable to vary across jobs. Thus, our level-2 model is:

$$\pi_{0jk} = \beta_{00k} + \beta_{01k}(\text{job \%female}_{jk}) + \beta_{02k}(\text{job \%female}^2{}_{jk}) + \beta_{03k}X_{1jk} + \ldots + \beta_{0qk}X_{qjk} + r_{0jk}$$

$$\pi_{1jk} = \beta_{10k} + \beta_{11k}(\text{job \%female}_{jk}) + \beta_{12k}(\text{job \%female}^2{}_{jk}) + \beta_{13k}X_{1jk} + \ldots + \beta_{1qk}X_{qjk} + r_{1jk}$$

where β_{00k} is the intercept for the job-level model in local industry k. In turn, β_{01k} and β_{02k} are the effects of job proportion female and its square on π_{0jk}. Likewise, β_{10k} is the job-level intercept for the effect of being female, π_{1jk}, while β_{11k} and β_{12k} are the effects of job proportion female and its square on the level-1 effect of being female (these are cross-level interaction effects). Finally, X_{1jk} through X_{qjk} denote the Q control variables in each job-level model. These control variables are centered around their grand means. The level-2 error terms are denoted by r_{0jk} and r_{1jk}. These error terms mean that the coefficients for the intercepts and job composition are within-local-industry effects, just as the intercept and gender coefficient in the individual-level model are within-job effects.

Finally, each level-2 coefficient relating job characteristics to level-1 effects on wages can be modeled as either a random or a fixed effect across local industries. In our models, only the level-2 intercept and the job proportion female (and its square) coefficients are permitted to vary across local industries. Thus, our level-3 model is:

$$\beta_{00k} = \gamma_{000} + \gamma_{001}(\text{mngr \%female}_k) + \gamma_{002}(\text{mngr ND}_k) + \gamma_{003}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{00s}W_{sk} + u_{00k}$$

$$\beta_{01k} = \gamma_{010} + \gamma_{011}(\text{mngr \%female}_k) + \gamma_{012}(\text{mngr ND}_k) + \gamma_{013}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{01s}W_{sk} + u_{01k}$$

$$\beta_{02k} = \gamma_{020} + \gamma_{021}(\text{mngr \%female}_k) + \gamma_{022}(\text{mngr ND}_k) + \gamma_{023}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{02s}W_{sk} + u_{02k}$$

$$\beta_{10k} = \gamma_{100} + \gamma_{101}(\text{mngr \%female}_k) + \gamma_{102}(\text{mngr ND}_k) + \gamma_{103}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{10s}W_{sk} + u_{10k}$$

$$\beta_{11k} = \gamma_{110} + \gamma_{111}(\text{mngr \%female}_k) + \gamma_{112}(\text{mngr ND}_k) + \gamma_{113}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{11s}W_{sk} + u_{11k}$$

$$\beta_{12k} = \gamma_{120} + \gamma_{121}(\text{mngr \%female}_k) + \gamma_{122}(\text{mngr ND}_k) + \gamma_{123}(\text{mngr \%female}_k \times \text{mngr ND}_k) + \ldots + \gamma_{12s}W_{sk} + u_{12k}$$

where γ_{000}, γ_{010}, γ_{020}, γ_{100}, γ_{110}, and γ_{120} are the level-3 intercepts in models of the level-2 coefficients; γ_{001}, γ_{011}, γ_{021}, γ_{101}, γ_{111}, and γ_{121} are the effects of the gender composition of managers on the level-2 coefficients. Likewise, γ_{002}, γ_{021}, γ_{022}, γ_{102}, γ_{112}, and γ_{122} denote the effects of the status of female managers (ND) on the level-2 coefficients. Finally, γ_{003}, γ_{013}, γ_{023}, γ_{103}, γ_{113}, and γ_{123} are manager percent female by manager relative status interaction terms. Coefficients for the *S* level-3 control variables *(W)* are denoted by the remaining γ terms. They are centered around their grand means. The level-3 error terms are given by u for each of the k local industries.

RESULTS

Results from the hierarchical linear model appear in Table 4. We show the coefficients only for key variables; complete results are available from the authors. The variance components for several models are presented in the Appendix.

Recall that in our models gender is associated with individuals' wages through three distinct pathways: individual gender, the gender of the co-workers in their jobs, and the gender of the

Table 4. Coefficients for Three-Level Hierarchical Linear Regressions for Logged Wages On Individual, Job, and Local-Industry Characteristics

Level 1 Individual Variables	Level 2 Job Variables	Level 3 Local-Industry Variables	Coefficient	t-statistic
INTERCEPT			2.888***	236.85
		Manager Percent Female	–.002***	–6.43
		Index of Net Difference	–.056	–1.22
		Manager Percent Female × Net Difference	.002	1.66
	Job Percent Female		.124*	2.27
		Manager Percent Female	.001	.52
		Index of Net Difference	.324	1.46
		Manager Percent Female × Net Difference	–.018**	–3.01
	Job Percent Female^2		–.385***	–6.00
		Manager Percent Female	.002	1.09
		Index of Net Difference	–.268	–.99
		Manager Percent Female × Net Difference	.017*	2.55
FEMALE			–.126***	–7.98
		Manager Percent Female	–.0003	–.68
		Index of Net Difference	.026	.40
		Manager Percent Female × Net Difference	–.002	–1.15
	Job Percent Female		–.270***	–3.72
		Manager Percent Female	.004*	2.09
		Index of Net Difference	–.456	–1.47
		Manager Percent female × Net Difference	.023**	2.72
	Job Percent Female^2		.327***	4.34
		Manager Percent Female	–.004*	–2.17
		Index of Net Difference	.475	1.46
		Manager Percent Female × Net Difference	–.023**	–2.85

Source: 2000 U.S. Census Public Use Microdata files and other sources (see text).
Notes: This model includes control variables at all three levels (see Table 1). Coefficients in the lower panel, on the individual female coefficient, are cross-level interactions showing effects on the wage difference between men and women.
* $p < .05$; ** $p < .01$; *** $p < .001$ (two-tailed tests).

managers in their local industries. At level 1, the gender of individuals is associated with wage differences within jobs. At level 2, the gender composition of jobs is associated with the average wages of men and women across those jobs within local industries. This effect may be different for women and men and may be nonlinear. At level 3, the gender characteristics of managers are associated with overall wage levels, as well as with the job- and individual-level effects. The variables at each level are labeled in three columns in Table 4. Because the number of interactions complicates the interpretation of the results, we only briefly discuss the coefficients before moving to a plot of predicted wages for men and women under plausible scenarios.

For men, the predicted wage at the mean of all control variables is the model intercept: 2.888 in logged dollars or \$17.96 per hour. This reflects an all-male job with no female managers and a net difference of zero because those variables are not centered. The individual-level gender effect is –.126, which means that women in the same job are predicted to earn 2.888 – .126 = 2.762, or \$15.83, for a gender wage ratio of .88. To extend the example simply, if 50 percent of the managers in that local industry are women, those managers are of equal status as the male managers (ND = 0), and the job remains all-male, then the predicted wage for men would be the model intercept plus the manager percent female effect: 2.888 + (–.002 × 50) = 2.788, or \$16.25. To calculate the predicted wage for women in such a job, we start with men's wage, then add the individual gender effect as modified by the manager composition of 50 percent female: 2.788 + [–.126 + (–.0003 × 50)] = 2.747, or \$14.11, for a gender wage ratio of .87.

This simple example is not realistic, however, because it reflects men and women working in an all-male job with equal status between male and female managers. When we assess the gender gap at different levels of job and managerial gender composition, and female managers' relative status, a more conclusive result emerges. Specifically, managerial composition and relative status largely work on the gender gap through the effects of job composition.[20] Note the significant effects of manager percent female and its interaction with ND on the level-2 variables in the lower half of the table. The net result is that *female managers are associated with a reduced gender wage gap* especially *when those female managers hold relatively high-status positions.* We will illustrate the results graphically rather than walk through a summary of many interactions and nonlinear effects.

Figure 2 shows predicted wages (the *y*-axis) for male and female workers in jobs of typical gender composition (30 percent female for men, 70 percent female for women), working under manager compositions ranging from 0 to 80 percent female (the *x*-axis), with relative gender status one standard deviation above and below the mean (one line for each). The predictions show men's wages are consistently lower where the percentage female among managers is higher (consistent with the suggestion that women managers are concentrated in less highly-paid industries). Among female workers, wages fall as a function of female management where those managers are of low status, but wages rise where female managers are of high status. For our focus on the gender wage gap, the result is clear. Net of controls at all three levels, the higher presence of female managers in local industries is associated with a reduced gender wage gap among nonmanagerial workers only where those managers hold relatively high-status positions. In the figure, the gender wage ratio is constant at .81, where female managers' relative status is one standard deviation below the mean, but narrows from .76 to .99 when that status is one standard deviation above the mean and manager composition rises to 80 percent, closing the gender gap.[21]

The results support both of our hypotheses. First, on average (i.e., midway between the high and low lines in Figure 2) the gender gap is smaller in local industries where a high pro-

[20] The baseline effect of job composition is such that, as female representation at the job level rises, men's wages start slightly upward but then turn sharply downward after 16 percent female (the inflection point is [–1 × .124]/[2 × –.385] = .16). The nonlinearity is much less pronounced for women, resulting in a negative net effect of female job composition.

[21] Such extreme cases are rare but do exist. In the sample, 90 percent of employees work under managerial pools that are between 12 and 67 percent female.

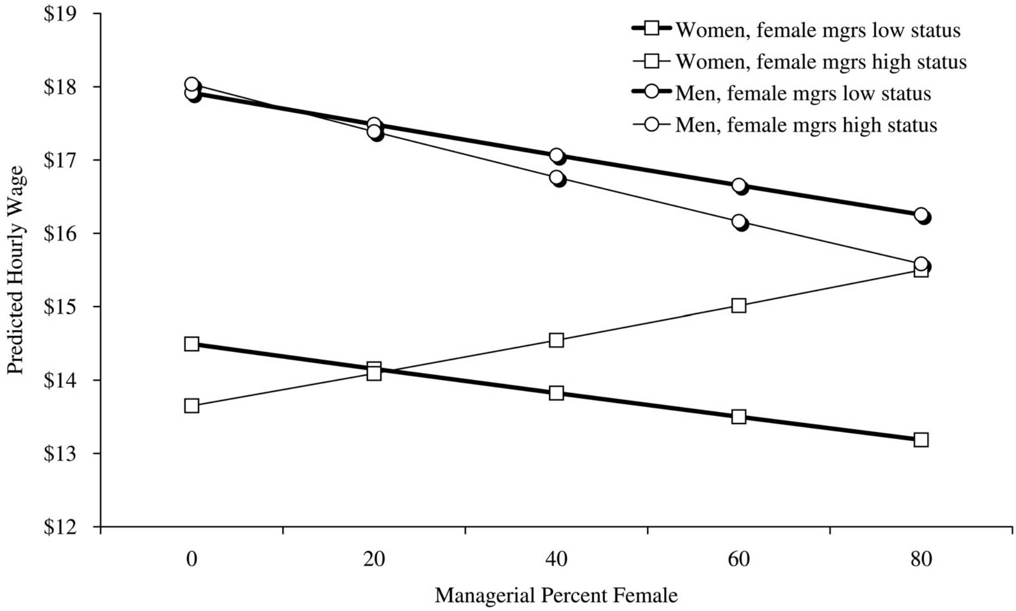

Figure 2. Predicted Wages for Men and Women, by Manager Percent Female and Relative Status

portion of the managers are women. This effect is small, though, and results primarily from a negative effect on average (men's) wages, net of female managers' relative status, women's wages are largely unaffected by manager percent female. The results also strongly support our second hypothesis, which posits an interaction between the representation and relative status of female managers. Specifically, we hypothesize less gender wage inequality in local industries with (1) more female managers and (2) female managers averaging higher-status positions. This is consistent with our data. *The gender wage gap is smaller under female managers, and this effect is much stronger when those female managers are of relatively high status.* Thus, although female managers may boost women's wages relative to men, the representation of low-status female managers alone does not affect the gender wage gap. It is representation in upper-status managerial positions that is associated with the gender wage gap.

ENDOGENEITY

Although the statistical relationship is strong, potential sources of error or bias exist in our analysis. In particular, endogeneity and the omission of potentially relevant control variables may be sources of bias, although our research design and choice of control variables help to minimize their effects. Endogeneity could confound our results to the extent that unmeasured variables are driving both the gender wage gap and the representation of women in management. However, we include controls for key indicators of gender inequality, including gender segregation and demand for female labor in the local labor market, as well as a proxy for the level of rationalization or bureaucratization of the local industry, which is likely to reduce the reliance on ascriptive characteristics in determining work-related rewards. Differences across local industries in endogenous factors that drive both the wage gap and female representation in management are likely to be captured by these controls. Controls at the job and individual levels play similar roles. In addition, because the job-level intercepts of our models are permitted to vary across local industries, and the individual-level intercepts vary across jobs, our analysis accounts for unobserved factors influencing average wages across local industries. In essence, then, our results reflect within-job and within-local-industry effects. This improves confidence in our results.

Alternative Specification

An additional problem could result from our definition of managerial authority. Identifying managers using occupational classifications is perhaps more straightforward than identifying those who are *not* managerial. Some people identified as nonmanagerial in our sample (e.g., doctors) may wield considerable authority, in some cases, more authority than those counted as managers (e.g., administrative service managers in doctors' offices). This represents the long-standing problem of intermediate class locations that preoccupies some analysts of class (e.g., Wright 1997). Excluding self-employed workers presumably helps reduce the number of such cases, but many workers in professional occupations remain who have ambiguous authority relations.

To verify that our results are not unduly influenced by such ambiguity, we reestimated our regression models but excluded from the nonmanagerial sample those in "business and financial operations" or "professional and related" occupations. This reduced the worker sample by more than one-third (39 percent of women and 29 percent of men). In fact, three of the largest occupations in our original sample were excluded: elementary and middle school teachers, registered nurses, and accountants and auditors (these three occupations alone accounted for almost 10 percent of the original sample). The largest remaining nonmanagerial occupations were unambiguously nonmanagerial: secretaries, truck drivers, and customer service representatives together compose 13 percent of the reduced sample. Results from this alternative specification (not shown, but available from the authors upon request) are substantively identical to those reported above. All the central coefficients are of the same or larger magnitude and remain statistically significant.

DISCUSSION AND CONCLUSIONS

After analyzing a survey of managers' attitudes nearly two decades ago, Brenner and colleagues (1989:668) concluded: "Unlike her male counterpart, today's female manager would be expected to treat men and women equally in selection, promotion, and placement decisions." Unfortunately, as they noted, women in management mostly held lower- and middle-management positions, where their impact was limited. Despite two decades of movement toward managerial integration, we still do not have conclusive evidence that the entry of some women into managerial positions has brought material benefits to the majority who remain below.

To address this question, we offer the first large-scale analysis using nationally representative data on workers tied to managers and including theoretically relevant variables at the level of the individual, job, and local industry. We consider the relative status of female managers as well as their numeric representation in models of the gender wage gap. Specifically, we test two hypotheses: (1) female workers earn more when their local industries include more women among the managerial ranks and (2) such representation is more beneficial when the relative status of female managers is higher.

Although we cannot draw causal conclusions from our data, our results are consistent with the argument that female managers do matter. In the models, the representation of women in management reduces the wage gap. The interaction between managerial gender composition and female managers' relative status, however, shows that the relationship is much stronger in local industries where female managers hold relatively high-status positions. The addition of women at the low end of the managerial hierarchy may have weak effects on the gender wage gap, but the potential effects of high-status female managers are much more positive.

This finding highlights—in a new way—the significance of the "glass ceiling." If our findings hold, not only are qualified women blocked from upper-level managerial positions and denied the benefits of those jobs, but their absence has ripple effects that shape workplace outcomes for nonmanagerial women as well. To the extent that women continue to cluster at the low end of managerial hierarchies, our findings may temper the optimism generated by the rapid increase in the proportion of women in management in the last several decades. We are also given pause by the finding that, with all the controls in the model, wages for men are lower in local industries with more female managers. This suggests that female managers remain concentrated in workplace settings with lower wages across the board, in ways that we cannot capture with the variables used here. On the other hand, our findings imply that inroads made by women into upper-status managerial positions will "lift all boats" by also boosting the wages of women employed in nonmanagerial occupations. All women may benefit from the desegregation of managerial occupations, even those who do not themselves attain such positions—which compliments Cotter and colleagues' (1997) finding that all women benefit from occupational desegregation.

In addition, our results are consistent with Jacobs's (1992) notion of "title inflation"—the reclassification of previously nonmanagerial workers as managers with little change in authority or wages—in that the mere representation of women in management has limited effects on the gender wage gap. Importantly, our analysis shows that simply looking at the percentage of females among managers is not enough if one is interested in the *effects* of managerial access. In this case, including the relative status of female managers is necessary—this assertion garners strong support from our multivariate results.

Our results are intrinsically sociological and broadly provocative. We hope they will entice others to investigate not only the role of managers, but also analogous cases of subordinate group members in positions of authority. For example, in studies of political alienation and participation, research shows that the presence of Black (Bobo and Gilliam 1990) and Latino (Pantoja and Segura 2003) elected leaders increases political empowerment among members of those groups. Female and minority political leaders appear to be more responsive to the concerns of their ascriptively similar constituents (Bratton and Haynie 1999; Mansbridge 1999), but the long-term implications of such integration remain to be seen. In education, students receive more positive evaluations and feedback from teachers who match their race/ethnicity and gender, but the evidence that this results in better educational outcomes is limited and mixed (Butler and Christensen 2003; Dee 2005; Downey and Pribesh 2004; Ehrenberg, Goldhaber, and Brewer 1995).[22] As with our examination of labor markets, these studies confront possible confounding effects across multiple levels of social interaction. Researchers in these disparate fields should learn from each other.

Further theorizing and data collection efforts undoubtedly will help to overcome the weaknesses in our case. For example, we are not able to link workers and managers in their actual work settings. Just as workplace processes generating inequality vary across organizational settings (Baron and Newman 1990; Huffman and Velasco 1997), female managers' ability to alter wage setting practices also may vary markedly across workplace contexts. On the other hand, in our analysis the presence and status of female managers is measured at the level of local industries, which may be the relevant organizational field within which employers make crucial gender-related decisions. This permits us to extend existing research on the gender wage gap to incorporate managerial characteristics using high-quality, large-scale Census data that links population-based samples of workers to managers. We hope this innovation, and these results, will generate increased attention to the role managers play in producing and sustaining labor market inequality, and by extension to the potential influence of subordinate group members who attain positions of authority.

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[22] Ehrenberg and colleagues (1995:548) note: "The relationships between supervisors and employees is analogous, in important respects, to that between teachers and students." They conclude, "Research addressing these issues should be high on the priority list of those concerned with the progress of women and minorities in the labor market" (p. 560).

gender and racial differences in access to managerial jobs, and the consequences of changing managerial composition for inequality among workers.

APPENDIX

Variance Components

The hierarchical linear models permit us to decompose the total variance in wages into three component parts: (1) that which occurs between individuals and the mean wages in their jobs, (2) that which occurs between mean wages in each job and the mean wages for all jobs in their local industries, and (3) that which occurs between the mean wages for local industries. Results from the unconditional model (which includes no independent variables at any level) in Table A show that overall, 72.3 percent of wage variation among nonmanagerial workers occurs within jobs as we have defined them; 14.8 percent occurs between jobs but within local industries; and 12.9 percent occurs across local industries.

This construction differs from previous three-level wage decompositions (e.g., Cohen and Huffman 2003a) in its use of local industries as the third level rather than metropolitan labor markets. Not surprisingly, variation across local industries accounts for a larger share of the total variation than does metropolitan-area labor markets. The differences between average wages for the restaurant industry versus the computer systems industry within a given labor market would be expected to be larger than differences in average wages between New York and Los Angeles overall. The variance components also show that our variables do much more to explain variance between jobs and local industries (variance components are reduced by about 80 percent each from the unconditional model to the final model) than they do to explain variation between individuals, which is only reduced 8 percent in the full model. Finally, Table A shows that about three-fourths of the variance in the effect of being female occurs across jobs but within industries. In the final model, 42 percent of that within-local-industry variance is explained, as the variance component is reduced from .0104 to .0068.

Table A. Variance Components for Three-Level Hierarchical Linear Regressions

	Unconditional Model (UM)		UM + Female Dummy		Full Model	
	Variance Component	Percent of Total	Variance Component	Percent of Total	Variance Component	Percent of Total
Intercept						
Individual	.3459	72.3	.3404	71.7	.3191	92.3
Job	.0708	14.8	.0718	15.1	.0146	4.2
Local Industry	.0620	12.9	.0623	13.1	.0118	3.4
Total		100.0		100.0		100.0
Female						
Job			.0104	75.1	.0068	56.9
Local Industry			.0035	24.9	.0051	43.1
Total				100.0		100.0
Job Percent Female (on intercept)						
Local Industry					.0776	100.0
Job Percent Female Squared (on intercept)						
Local Industry					.1031	100.0
Job Percent Female (on individual female effect)						
Local Industry					.1055	100.0
Job Percent Female Squared (on individual female effect)						
Local Industry					.1051	100.0

Note: All variance components have a two-tailed p-value of less than .05.

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