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Occupational Segregation and the Gender Gap in Workplace Authority: National Versus Local Labor Markets

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Previous research linking occupational gender segregation to the workplace authority gap assumes that the effect of gender composition is invariant across occupations, ignoring the important distinction of whether an occupation's relevant labor market is local or national. We offer a new method for defining occupational labor markets and hypothesize that the effect of occupation gender composition on the authority gap will be strongest in national labor market occupations. Both sexes' odds of possessing work authority decline with the representation of women; this effect is strongest in the more desirable, national labor market occupations. Assuming occupations are part of one labor market results in understating the gender composition penalty for national labor market occupations.

KEY WORDS: occupational gender segregation; workplace authority; labor markets.

The capacity to exercise authority over organizational resources and coworkers is central to workers' overall work experience (Halaby, 1979; Reskin and Padavic, 1994; Reskin and Ross, 1992). Wright *et al.* (1995) note that authority is a highly valued attribute of jobs because it is status-conferring and shapes how financial rewards are allocated to workers. Furthermore, they argue that the paucity of women occupying authority positions in the workplace is not merely an instance of gender inequality, but a significant cause of gender disparities as well (407). The importance of legitimate control over resources and workplace power for understanding

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work-based inequalities has also been shown by many others scholars (e.g., Halaby, 1979; Jaffee, 1989; Spaeth, 1985; Wallace *et al.*, 1993; Wolf and Fligstein, 1979; Wright, 1979).

Do female-dominated occupations offer fewer opportunities for workplace authority? In this paper, we address the gender gap in work-based authority, paying special attention to how it relates to occupational gender composition. We also consider another important point: whether differences in *occupational labor markets* influence the association between gender, gender composition, and workplace authority. Although prior research has ignored variation in the relevant labor markets for different occupations, it is important because the arena in which competition over scarce resources such as workplace authority takes place may be very different for high-status occupations, which operate in a national labor market, and for occupations that operate in more circumscribed, local labor markets. To examine this relationship, we respecify the effect of occupation gender composition by allowing it to vary across occupations in national and local labor markets. This innovation allows us to show several important and unexpected features of the relationship between gender, gender composition, and workplace authority.

We make several distinct contributions to the existing literature on gender-based workplace inequality and the interaction between labor markets, gender, and workplace rewards. First, we employ a unique method for capturing the local/national dynamics of occupations, on the basis of 1990 Decennial Census data and data from 261 U.S. metropolitan areas. We then use this operationalization of labor markets to provide strong empirical evidence that the gender composition of an occupation has a much larger effect on the authority levels of its members if the occupation operates in the national labor market rather than in local labor markets.

WHO'S THE BOSS? THE GENDER GAP IN WORKPLACE AUTHORITY

A large body of research documents the underrepresentation of women in positions of authority at work (Jaffee, 1989; Kanter, 1977; Reskin and Ross, 1992; Wolf and Fligstein, 1979; Wright and Baxter, 2000). If the gender authority gap primarily reflects men's desire to protect their advantaged positions, there is scant reason to believe that declining human capital differences between women and men will produce equality in the distribution of authority (Reskin, 1988; Reskin and Ross, 1992). Moreover, shrinking gender differences in job-related skills and aspirations may not reduce employer discrimination that is rooted in the belief that women's emotions prevent them from managing effectively (Kanter, 1977) or in normative workplace standards that preclude women from exercising power over men Bergmann, 1986; Kanter, 1977; Schroedel, 1985).

As in the case of wage inequality, gender differences in human capital account for only part of the workplace authority gap, leaving much of it unexplained. For example, Huffman (1995) finds that women are less likely to possess supervisory authority at work than men with equivalent levels of education, occupational experience and prestige, and family characteristics such as marital status and the presence of children. Jaffee (1989) reports similar findings: Across three indicators of workplace authority and autonomy, women suffer a large disadvantage relative to men, controlling for variables stressed by supply-side explanations. Not only is there a large gender gap in work-based authority, but women also accrue additional disadvantage through lower earnings returns to authority (see Reskin and Ross, 1992; Wolf and Fligstein, 1979). Substantial authority differences appear also in cross-national research (see Rosenfeld *et al.*, 1998; Wright *et al.*, 1995; Wright and Baxter, 2000).

Thus, empirical findings describing the gender authority gap are at odds with supply-side accounts, which stress gender differences in the household division of labor to explain disparities across an array of occupational outcomes (e.g., Becker, 1985; Polachek, 1979). According to this perspective, the gender gap in authority results from gender differences in the allocation of time and effort to various tasks, as well as differences in the kinds of investments made in productive skills (Jaffee, 1989). Because of numerous exogenous factors, women are purported to have a competitive advantage in the performance of household duties, while men specialize in paid labor. Because of this, women self-select into female-dominated occupations, and put forth less effort than men in the workplace. This rational decision-making process occurring at the household level makes women less likely to be promoted into positions of authority than similarly qualified men. (For a critique, see Bielby and Bielby, 1988.) Gender specialization at the household level has also been used to explain other outcomes, such as the tendency for married men to earn more than single men (e.g., Hersch and Stratton, 2000).

Not only does the gender authority gap persist after making adjustments for human capital, but findings regarding the interaction of family structure variables and gender are largely inconsistent with predictions based on presumed occupational self-selection. Those explanations posit that the presence of children in the household induces women not to pursue jobs involving authority, while men are largely unaffected by children (Wright *et al.*, 1995). However, several studies demonstrate that marriage and the presence of children are only weakly associated with the gender composition of the jobs people hold. This contradicts the argument that women forgo promotion opportunities and other rewards for jobs offering work–family compatibility (see Glass and Camarigg, 1992; Jacobs, 1989; Okamoto and England, 1999).

Occupational Gender Segregation and Workplace Authority

Much attention has been devoted to documenting the effect of the gender mix of jobs and occupations on diverse outcomes, including the gender earnings gap (Baron and Newman, 1989; England, 1992; Petersen and Morgan, 1995; Reid, 1998; Tam, 1997), job satisfaction and psychological orientations toward work (Wharton and Baron, 1987, 1991), and gender stereotyping and evaluation bias (Konrad *et al.*, 1992; Tsui and O'Reilly, 1989). However, only sparse research has addressed gender composition effects on the authority gap, even though occupational gender segregation is a main avenue for discrimination in access to authority (Jaffee, 1989; Kraus and Yonay, 2000). One reason that women lack authority is that most women are concentrated in female-dominated occupations, which comprise fewer positions of authority than male-dominated occupations. Indeed, Huffman (1995) finds the concentration of women in predominantly female occupations to be a strong predictor of women's authority deficit, even after making adjustments for worker characteristics. Specifically, after controlling for supply-side factors, nearly two-thirds of the remaining gender difference in authority was accounted for by occupational-level gender composition. Jaffee (1989) also reported a substantial disadvantage for women when only human capital variables (and not gender composition) were controlled; this, and the finding that most of the gender gap is explained by occupation gender composition, suggest that the nature of the job is more important than individual-level factors in explaining women's authority deficit. If women and men with similar human capital profiles are channeled into occupations that differ in terms of the number of available authority positions, occupation gender composition would clearly dwarf human capital effects.

What is the mechanism producing the association between gender composition and job rewards such as authority? One possible answer to how the representation of subordinate groups shapes inequality is found in the "competition" (or "visibility-discrimination") hypothesis, which derives from the work of Blalock (1967), although it has roots in earlier social psychological research by Allport (1954). At any time, a fixed number of available jobs exist in the labor market. The competition hypothesis predicts more discrimination when a subordinate group is larger, because of the heightened competition over scarce resources. Although the competition hypothesis was originally applied to race at the community level, it has also been tested at the labor-market level (e.g., Beggs *et al.*, 1997; Cohen, 2001; Jacobs and Blair-Loy, 1996; Tienda and Lii, 1987). As Reskin *et al.* (1999:345) note, the competition hypothesis applies to gender composition as well.

With respect to the gender gap in workplace authority, some have used the competition hypothesis to predict that increases in the representation of women in an *occupation* will intensify competition over authority positions and consequently magnify gender discrimination (Kraus and Yonay, 2000). This implies that the largest authority gap will exist where women are the majority. The gender gap should be relatively small in male-dominated occupations, because gender-based competition is reduced. In their test of this prediction, Kraus and Yonay (2000) find a markedly smaller gender authority gap in male-dominated occupations than in female-dominated occupations, supporting the competition hypothesis.

In contrast, one might predict that there will be less gender discrimination in access to authority positions where women predominate. Kanter (1977) was among the first to investigate how women's proportional representation influences their work experiences. Her work supports the contention that increased representation of women should reduce discrimination. Similarly, Blau (1977) posited that as heterogeneity increases, group membership becomes less significant to the "in-group," and, as a result, their propensity to discriminate declines. Although female-dominated occupations comprise fewer positions of authority than male-dominated occupations, women's presence may counterbalance their disadvantage by increasing their power and leverage (Kraus and Yonay, 2000). Ely (1994:209) hypothesizes that women employed in firms with higher proportions of upper-level women may experience a more hospitable working environment, which may ease struggles over workplace resources. However, some researchers characterize as overly optimistic the view that gender-balance in an occupation or job will eliminate the negative effects of tokenism suffered by women (see Ely, 1995; Yoder, 1991). Reskin *et al.* (1999:345) point out that Blau's more recent work (see Blau, 1994) acknowledged that the effect of gender is strong enough to overcome the structural effects of relative group size.

Thus, there are two ways that increasing representation of women might promote gender equality. One points to social psychological processes through which public images and gender stereotypes are modified in ways that benefit women as their representation increases. The other highlights the power that may accompany the presence of women employed in a particular occupation or work setting. However, a third alternative to the competing explanations derived from the work of Blalock/Allport and Kanter/Blau is that men are uniformly advantaged across occupations, regardless of the occupation gender composition. For example, Budig (2002) finds that men's wage advantage is constant across male-dominated, gender-mixed, and female-dominated occupation-industry cells.

Gender Composition Effects Across Occupational Labor Markets

Researchers interested in issues surrounding workplace inequality often specify statistical models that assume a single U.S. labor market; however, many labor markets make up the U.S. economy. Workers in high-skill occupations compete for jobs in a national labor market, while unskilled or semiskilled workers compete in local labor markets. In those occupations, workers seek employment, and are recruited by employers, within a reasonable commuting distance from their homes (Baker, 2000; Helfgott, 1974).

Why is the labor market scope of an occupation relevant for questions about the gender authority gap? We argue that whether an occupation is embedded in a national versus a local labor market is integral to how competition is conceptualized, and therefore an analysis that makes such a distinction is better specified. For example, the presence of underrepresented groups in a broad occupational category may indeed increase competition (and therefore discrimination), but only among those workers who compete in the national labor market. In this case, increases in the representation of women may be assumed to taint the prestige of the occupation (Reskin and Roos, 1990). In contrast, for those in occupations that compete in local labor markets, the salience of broad, macro-level changes in demographic composition is muted. For example, a male brickmason might not be concerned about an increase in the proportion of female brickmasons nationally, but he might react strongly to the presence of female brickmasons at his work site. This is because brickmasons compete in a local market, making changes at the broad occupational level of secondary concern. The fact that there are increasing numbers of female brickmasons nationally does not imply that there will be heightened competition with women in a particular local labor market. So, to the male bricklayer in Seattle, the number of female bricklayers in Los Angeles will be of secondary concern, because he is not in direct competition with these women. However, a male engineer is likely to object to shifts in the gender composition of his occupation, even if he never encounters a female engineer at his workplace. This is because this male engineer works in a national market and must therefore compete with all female engineers entering the profession, regardless of geographic location. Thus, a male engineer in Seattle will care if women are increasingly becoming engineers in Los Angeles, because he will be forced to compete with them, since the labor market for engineers is national.[3]

Metaphorically, the site where competition is played out expands or contracts depending on what kind of occupation is being considered—when the work is relatively high-status, the playing field may be properly specified at the national occupational level. However, competition over relatively low-status work, because of its organization based on local labor markets, may be less likely to be altered by changes in the demographic composition occurring at the level of national occupations. In other words, measuring the gender composition of occupations at the national level may be appropriate for occupations that operate in the national labor market but not for those that operate in local labor markets. Despite this, previous research has ignored this source of variation in the gender composition effect on authority. For example, Huffman (1995), Jaffee (1989), and Kraus and Yonay (2000) all model the effect of gender composition as invariant across occupations.

RESEARCH AGENDA

We address four issues regarding the gender gap in workplace authority. First, we document the extent of the gender authority gap and directly address supply-side explanations for the gap, which emphasize the self-selection of women into low-authority positions because of the presence of children and other family responsibilities.

Second, we address the size of the gender gap in authority across levels of occupation gender composition, without distinguishing between occupational labor markets. We then consider the type of labor market (national versus local), asking whether the effect of occupation gender composition is particularly strong among occupations in the national labor market, meaning that women's presence in an occupation lowers the prevalence of authority positions more than in occupations embedded in local labor markets. This question does not address the size of the gender gap within occupations

[3]Even though a substantial proportion of those in managerial–professional occupations have workplace authority (as we have measured it), many do not. With respect to gender differences, the data we analyze (the 1991 General Social Survey) show that within managerial–professional occupations, 62% of men have authority, compared with only 46% of women. Thus, although the meaning of authority may differ between occupations in the local and national labor markets (for example, those with authority in managerial–professional occupations may supervize a greater number of subordinates), there is plenty of room for gender-based struggle over work-based authority in those occupations that tend to be in the national labor market. This struggle might result in job-level gender segregation—as women tend to enter occupations in the national labor market, they are assigned to jobs with low levels of authority (see Jacobs, 1992).

with varying gender composition. Instead, it asks whether the finding that female-dominated occupations comprise fewer positions of authority holds uniformly across occupations in national versus local labor markets. An affirmative answer to this general question of whether there are fewer positions of authority in female-dominated occupations suggests bias against lines of work typically performed by women. This would parallel empirical findings on the effect of occupation gender composition on earnings (e.g., England, 1992).

The last portion of the analysis considers whether the effect of gender composition on the size of the gender authority gap varies by labor market scope (national versus local). This implies an interaction between respondents' gender, occupation gender composition, and occupational labor market. Here, we address the competition hypothesis, while increasing its specificity. We hypothesize that occupational-level changes in gender composition affect levels of competition most strongly in occupations in the national labor market—thus, the largest gender gap in authority should be found in heavily female occupations, but this gap should be larger in occupations in the national labor market than in those in local labor markets. We address this question by examining how the size of the gender authority gap covaries with occupation gender composition, in models estimated separately for national versus local labor markets.

DATA, MEASURES, AND MODELS

Data Sources

Our primary data source is the 1991 General Social Survey (GSS). The GSS is an annual personal interview survey of a probability sample of English-speaking U.S. adults conducted by the National Opinion Research Center. In 1991 the GSS included a special topical module on work organizations that included a series of questions about workplace authority and other job-related duties, satisfaction with work, and other items pertaining to employment.

The measure of occupation gender composition comes from a second data source, the Census Bureau's 1990 Equal Employment Opportunity (EEO) Supplemental Tabulations File (Census of Population and Housing, 1990). The GSS respondents' three-digit census occupational code was used to append the gender composition measure to each individual GSS record. We exclude self-employed respondents, as they are not subject to the decision-making of an employer. Listwise deletion of missing cases resulted in 876 cases to be used in the analyses.

Measures: Dependent Variable

Workplace Authority

The GSS respondents were asked, "As an official part of your job, do you supervise the work of others or tell other employees what work to do?" Affirmative responses were coded 1; those answering "no" were coded 0. This is the dependent variable in the multivariate models.

As noted by Wright *et al.* (1995:417), this definition of authority is largely independent of the occupation in which a person works, because many individuals employed in lower white-collar occupations have substantial amounts of work-based authority.[4] Thus, the measure used here directly addresses whether one has legitimate authority as part of his or her job. Similar dichotomous indicators have been used in previous research on the gender authority gap (e.g., Adler, 1993; Jaffee, 1989; Kraus and Yonay, 2000; Wright *et al.*, 1995). Additionally, the dichotomous indicator used here provides a conservative estimate of the gender authority gap: Among those men and women who possess workplace authority, men tend to occupy higher positions in the authority hierarchy than women (McGuire and Reskin, 1993). Thus, this measure provides a lower bound on the estimated gender authority gap.

Measures: Independent Variables

Gender Composition

As noted above, the measure of occupation gender composition comes from census tabulations. It is the percentage of females in each GSS respondent's three-digit occupational category.

Occupational Labor Markets

Our analysis requires a method for classifying occupations as either national or local in scope. An occupation can be said to function in the national labor market to the extent that competition over available jobs among workers (and competition among employers over workers to fill job openings) in the occupation is not spatially bound; for example, university professors

[4]In the 1991 GSS, 54% of those in managerial–professional occupations have authority, compared to 30% of those in other occupations.

compete in the national labor market because a large proportion of job seekers apply for jobs that are not geographically proximate. Similarly, universities advertise their open faculty positions in national publications and cast their net over the entire country when seeking qualified candidates. In contrast, workers in relatively low-skilled occupations (and employers seeking workers for such positions) compete in locally circumscribed markets. For example, those competing for an available janitorial position in Seattle compete only with other janitors from Seattle. Similarly, an employer in Seattle who wants to hire a janitor to would be unlikely to seek out candidates in Houston. We should note that the national versus local distinction is largely conceptual—few, if any, occupations have fully developed national labor markets, in the sense that most workers in an occupation consider the entire range of jobs when applying, and most employers advertise in national publications when seeking out workers. For example, academics employed by research universities have a highly developed national labor market; however, those in other kinds of universities commonly operate in more localized labor markets. Thus, there is a potentially important gap between our conceptual distinction and the reality of finding work (and workers). With this in mind, we describe our measure.

There is no available measure of labor market scope for three-digit occupational categories. However, labor economists (e.g., Ehrenberg and Smith, 1996) note that there is a positive relationship between an occupation's educational requirements and the probability that it operates in a national labor market.[5] As one ascends the occupational hierarchy, the relevant labor market expands from local to national, with most unskilled and semiskilled workers seeking jobs within a reasonable commuting distance from their residence, and high-skilled workers (including many professional, executive, and managerial workers) competing in the national market (Kaufman, 1991). Although education is one important dimension, we develop a measure that relies on a direct assessment of the local/national dynamics of occupations in conjunction with educational information.

Specifically, we use the differential effects of the two kinds of labor markets as symptoms that can identify their presence. These effects concern the composition of the labor force and how it varies—or does not vary—across

[5]Some research has examined whether neighboring metropolitan areas are part of the same labor market by examining the similarities in changes in pay rates for occupations common to any two adjacent metropolitan areas (e.g., Barkume, 1996). A statistically significant correlation between pay rates indicates interdependence in occupational wage developments in adjacent areas, and suggests that the two metropolitan areas are part of the same labor market. However, this method addresses whether two geographic regions are linked by a common labor market. This is not the question we target with our measure.

localities. If an occupation comprises workers embedded in local labor markets, then the compositional characteristics of the individuals in that occupation can be expected to reflect the overall composition characteristics of local areas. In contrast, if an occupation is national in scope, then the compositional characteristics of its incumbents can be expected to have less similarity to the local compositional characteristics (or greater homogeneity across local areas).

Thus, our method for defining occupational labor markets relies on the bivariate relationship between the *local* gender composition of the occupation, and the *local* degree of occupational gender integration. We employ the 1990 metropolitan area data file constructed by Cotter *et al.* (1997). This data set includes an occupational gender segregation index for each U.S. metropolitan area. The segregation measure is an adjusted dissimilarity index, which indicates the proportion of women who would have to change occupations so that the observed number of women in each occupation was no larger or smaller than chance would predict. We then use the 1990 U.S. Census (5% Public Use Microsample) data to calculate the percent female in each occupation for each of the 261 metropolitan areas in 1990. *If the local gender composition of an occupation is significantly correlated (in the absolute sense) with the degree of local occupational gender segregation, then it cannot be said the occupation operates independently of the local labor market—it is in a local labor market.* On the other hand, if the local gender composition of the labor market is largely independent of the local degree of occupational gender segregation, this suggests a national labor market may be operating.

For example, secretaries are approximately 99% female, nationally. However, the degree to which the local composition of secretaries is female-dominated depends on the level of local gender segregation, measured at the metropolitan-area level. For secretaries, the correlation is .29 ($p < 0.001$), indicating that in more segregated labor markets, the pool of secretaries is substantially more female-dominated. This is evidence that a local labor market is operating. The same is true of truck drivers, who are 6.1% female nationally. The local composition of truck drivers shows a statistically significant correlation with the local level of gender segregation—women are a markedly smaller share of the truck drivers in labor markets that are more segregated. In contrast, consider airplane pilots, of whom only 3.2% are women nationally. The correlation between the local composition of airplane pilots and the local level of gender segregation is not statistically significant ($r = -0.09$, $p = 0.37$), suggesting the labor market for pilots is national, not local. We should note that this dimension of our definition hinges on a significant *absolute* correlation—the direction of the association is not relevant.

As a further check, we also computed our measure using the correlation between percent female in the occupation locally (in the metropolitan area), and percent female (in the prime working ages of 16–64) in the metropolitan area. The two measures classify approximately 91% of the occupations in the same way ($\chi^2 = 211$, $p < 0.0001$). Using the measure based on local gender composition (rather than segregation) does not change our main empirical results substantially. We report results from models using local gender segregation.

Occupations displaying a statistically significant absolute bivariate correlation may be said to have strong local dynamics, enough to be counted as local, rather than national, labor markets. However, the *absence* of a statistically significant correlation does not necessarily imply the operation of a national labor market. Consider, for example, hairdressers/cosmetologists and construction laborers. Hairdressers/cosmetologists are 92.9% female nationally, and the level of local gender segregation and the gender composition of this occupation are essentially independent ($r = 0.002$, $p = 0.97$). Similarly, construction laborers are only 3.7% female nationally, and they show no correlation between local gender composition and local gender segregation ($r = 0.02$, $p = 0.75$). If we only applied the correlation rule, these occupations would be classified as operating in a national labor market. However, we would argue that these occupations are so firmly gender-typed that their gender composition is impervious to the variation in gender dynamics across U.S. labor markets. Importantly, the incumbents in these occupations have relatively low levels of education and earnings—these factors clearly do not justify national job searches on either the supply or the demand side. The correlation measure resulted in 12% of occupations being classified as national labor market occupations.[6]

Therefore, we apply a second rule, which relies on occupations' educational distribution. For occupations that *do not* show the correlation between local gender composition and local labor market segregation, we classify only those in which a large percentage of the incumbents nationally have college degrees as operating in national labor markets. Specifically, we use data from the 1990 U.S. Census (5% Public Use Microsample) to calculate the percentage of full-time, year-round workers (age 18–65) with a bachelor's degree or higher in each occupation (using the Bureau of Labor Statistics' definition

[6] As with any categorical measure, our strategy could result in a misclassification of occupations. We cannot rule out this possibility. However, if some occupations are misclassified, this would cause us to generally *understate* differences in the effect of our independent variables across labor market types. Additionally, one could argue that the labor market scope of occupations varies continuously from local to national. Further work, which we are currently performing, will be required to develop a continuous measure that allows fine-grained distinctions between occupations.

of FTYR employment).[7] We use the mean proportion among managerial-professional occupations (52%) as the cutoff.

We note that it is theoretically possible that a low correlation between local gender segregation and the local gender composition of an occupation could result from random variation in occupation gender composition across local areas, rather than labor market dynamics. In the presence of random (and substantial) variation, we might erroneously call an occupation national. Although it is possible, because we use educational levels in conjunction with the bivariate correlation, this potential source of bias could only apply to the 18% of our occupations that have "high" levels of education. Although we have no way knowing how many of this subset of occupations are misclassified due to random variation, our definition of labor markets—coupled with the fact that no previous research documents substantial random variation in gender composition—minimizes this possibility. Appendix A lists the five most male- and female-dominated occupations, by occupational labor market.[8]

Respondent Characteristics/Human Capital. The multivariate models include dummy variables for *gender* (female = 1; male = 0), *full-time* work status (1 = yes), and the *presence of children* (less than 18 years old) in the household (1 = yes). Human capital theory predicts that employers may use marriage to signal responsibility and stability among men but not women

[7]The 1991 GSS uses the 1980 Census codes to classify occupations, while the 1990 Census data uses the 1990 codes. However, the close correspondence between the two sets of codes allowed them to be easily matched. For example, the 1980 category "Managers and Administrators, not elsewhere classified (019)" was split into three 1990 codes, "Managers, Food Serving and Lodging Establishments (017)," "Managers, Services Organizations, not elsewhere classified (021)," and "Managers and Administrators, not elsewhere classified (022)." In this case, we assign the mean proportion with a college degree computed across the three 1990 codes to the 1980 code. Only one other 1980 code was split into multiple 1990 categories. It was coded using the same method. Additionally, several sets of two 1980 categories were each merged into a single 1990 category. For example, "Telegraphers (349)" and "Communications equipment operators, not elsewhere classified (353)" were merged into the single 1990 category "Communications equipment operators, not elsewhere classified (353)." In these cases, the proportion with a college degree computed for the 1990 category is assigned to both 1980 categories. This merging only applied to 12 of the 1980 categories (Six pairs). All of the other differences between the 1980 and 1990 codes involved either the title changing but the code staying the same (e.g. "Inhalation Therapists" became "Respiratory Therapists," and "Printing Machine Operators" became "Printing Press Operators"), or a change in numerical code but no change in title. These changes introduced no difficulty into translating between the 1980 and 1990 coding scheme.

[8]We considered potential problems arising from the use of an individual-level dependent variable and the way occupational labor markets are defined. For example, one might assume that our definition of occupational labor markets implies that those employed in national labor market occupations have a higher likelihood of having job authority than those in local labor market occupations. However, the zero-order relationship between the odds of having authority on the job and labor market scope (local versus national) is not statistically significant ($p > 0.10$).

(Becker, 1975; Tharenou, 1999). Therefore, we use a dummy variable for *marital status* (1 = married). The children and marriage dummy variables are interacted with sex to test whether they have sex-specific effects on the odds of possessing supervisory authority. We also include a continuous measure of *education* (highest year of schooling completed). Finally, to minimize gender differences in the accumulation of firm-specific human capital (Becker, 1975; Tam, 1997), a continuous measure of *tenure with current employer* (number of years working for present employer) is used.

Industrial Sector. Stratification by industry may affect women's mobility into management (Blum *et al.*, 1994:248). To account for differences in the distribution of opportunities to possess authority across industrial sectors, we use 10 dummy variables to represent the 11 industry categories represented in the Office of Management and Budget's Standard Industrial Classification. The categories are Agriculture, Construction, Nondurable Manufacturing (the omitted category in the multivariate analyses), 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. Table I reports descriptive statistics for all variables.

Table I. Descriptive Statistics for Variables used in the Analysis ($N = 876$)

Variable	Mean	SD	Range
Work authority and respondent characteristics			
Authority	0.36	0.48	0–1
Gender	0.51	0.50	0–1
Tenure with employer	13.66	16.50	0–54
Young children in household	0.43	0.50	0–1
Marital status	0.55	0.50	0–1
Education	13.59	2.72	3–20
Full-time work status	0.79	0.40	0–1
Occupation gender composition			
Occupation percent female	47.73	30.28	0.88–98.69
Occupational labor market			
National labor market	0.12	0.32	0–1
Industry			
Agriculture, forestry, and fisheries	0.03	0.18	0–1
Construction	0.07	0.25	0–1
Manufacturing (nondurable goods)	0.07	0.25	0–1
Manufacturing (durable goods)	0.10	0.31	0–1
Transportation, communication, and public utilities	0.07	0.25	0–1
Wholesale trade	0.03	0.18	0–1
Retail trade	0.14	0.34	0–1
Finance, insurance, and real estate	0.07	0.25	0–1
Business, repair, and personal services	0.11	0.31	0–1
Professional and related services	0.25	0.43	0–1
Public administration	0.06	0.24	0–1

Statistical Models and Analysis

Because the outcome variable is dichotomous, we use logistic regression models (Agresti, 1996). The coefficient on the gender composition variable is of principal interest because it suggests whether the representation of women in an occupation influences the likelihood that a respondent possesses work authority. For example, a negative effect of occupation percent female means that female-dominated occupations offer fewer opportunities to exercise authority than male-dominated occupations. The coefficient on the gender dummy variable (female = 1) is of equal interest; it represents the estimated gender gap in the likelihood of possessing authority. With no controls, the gender dummy variable yields the total gender gap. Models including the human capital and industry measures yield the net gap—the predicted gender difference in the likelihood of having authority between men and women with similar human capital profiles, employed in the same industry. The industry coefficients are omitted from the tables (they are available upon request).

The logistic regression model expresses the log-odds of possessing authority at work (on the left of the equal sign) as a function of the independent variables. The total gender gap is given by:

$$\mathrm{Log}[P/1 - \mathrm{P}] = \alpha + \beta_1 X_1 \quad (1)$$

where P is the probability of a respondent possessing authority, X_1 is the gender dummy variable (female = 1), and α is the intercept (men's predicted log-odds of possessing authority). The net gender gap is estimated by

$$\mathrm{Log}[P/1 - P] = \alpha + \beta_1 X_1 + \Sigma\beta_j X_j \quad (2)$$

where the specification is the same as in equation (1), except for the addition of X_j, a set of variables measuring respondent human capital, industrial sector, and gender composition. Finally, β_j are the corresponding partial regression coefficients. To facilitate interpretation of the results, the logistic regression coefficients are reported in an odds ratio metric in the tables. Odds ratios, which are interpreted as the multiplicative effect of the independent variable on the odds that the dependent variable takes a value of one, are obtained by exponentiating the logistic regression coefficients ($e^{\beta j}$)—odds ratios less than one indicate a negative relationship, while odds ratios greater than one signify a positive relationship (Allison, 1999).

To test whether the effects of the supply-side variables vary by gender, Gender × Marital status and Gender × Children interaction terms are added to equation (2). To test whether the size of the gender gap varies across levels of Gender composition, a Gender × Gender composition interaction term is used. To assess whether the effect of gender composition varies across national and local labor markets, a Labor market × Gender composition interaction term is included. Lastly, we examine the three-way interaction of gender, occupation gender composition, and occupational labor market, in order to address how the size of the gender gap varies by both gender composition and occupational labor market. For example, some (Kraus and Yonay, 2000) assert that Blalock's (1967) competition hypothesis implies a negative interaction between gender and occupation percent female, suggesting a larger gender gap in female-dominated occupations. We examine whether this interaction is stronger in the national labor market by testing the Gender × Gender composition interaction separately for occupations in each labor market type.

RESULTS

Gender, Gender Composition, and Workplace Authority

The first set of regression results appears in Table II. The first model includes only the gender dummy variable (female = 1), establishing the total gender gap in the likelihood of possessing workplace authority. Predictably, the total gender gap suggests that women have a significantly lower probability of possessing workplace authority than men. Specifically, the

Table II. Determinants of Gender Gap in Workplace Authority: Odds Ratios From Logistic Regression Models ($N = 876$)

Variable	Model 1	Model 2	Model 3	Model 4	Model 5
Main Effects					
Gender (female = 1)	.462***	0.560***	0.806	0.807	0.812
Education	—	1.15***	1.147***	1.148***	1.148***
Full-time	—	2.094***	1.956***	1.948***	1.990***
Tenure with employer	—	0.995	0.994	0.994	0.994
Married	—	1.287	1.226	1.203	1.252
Children	—	0.790	0.786	0.902	0.796
Occupation percent female	—	—	0.986***	0.986***	0.988***
National labor market (NLM)	—	—	—	—	2.740*
Interactions					
Gender × marital status	—	—	—	0.026	—
Gender × children	—	—	—	−0.036	—
Gender × occupation percent female	—	—	—	−.000	—
NLM × occupation percent female	—	—	—	—	0.978**
Control variables					
Industry controls included	No	Yes	Yes	Yes	Yes
Likelihood ratio chi-square[a]	29.7***	110.1***	116.4***	116.3***	121.3***

[a]Test of improvement in model fit versus the intercept-only model.
*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$ (two-tailed test).

odds of possessing authority for women are less than half as large (46.2%) as men's odds (OR (odds ratio) = .462). That is the zero-order gender effect.

Model 2 includes controls for respondent characteristics and industry, and allows one to see what fraction of the total gender difference is due to measured human capital differences among respondents. Predictably, the effects of education and full-time work status are positive and statistically significant (this holds across all models). Model 2 also shows a statistically significant gender effect; specifically, the finding indicates that even among women and men employed in the same industry who work comparable hours and have comparable levels of education, tenure with one's employer, and similar family obligations, women are markedly less likely to have authority at work. Among men and women who are comparable on those variables, women's odds of possessing authority are only 56% as large as men's odds (OR = .560). Thus, a large percentage of the gross gender gap is not attributable to differences in measured family obligations, differences in employment across industrial sectors, or other factors controlled in Model 2.

The third model includes occupation-level gender composition. The odds ratio for this variable (.986) indicates that for each percentage point increase in the percent female, both men's and women's odds of possessing workplace authority decrease by 1.4%, a statistically significant effect ($p < 0.001$). Thus, net of supply-side differences, all workers' odds of possessing authority are reduced when working in a female-dominated occupation. This effect mirrors previous work on gender earnings inequality (e.g., England, 1992), by showing a generalized gender effect whereby those occupations performed by women comprise fewer authority positions than typically male occupations. It is also noteworthy from Model 3 that the gender gap is no longer statistically significant when the gender composition variable is included. Comparing the gross gender gap in Model 1 with the net gender effect in Models 2 and 3 indicates that occupation percent female accounts for a much larger share of the gender authority gap than supply-side characteristics. This finding parallels research showing the stronger contribution of occupational gender segregation relative to supply-side differences in explaining the authority gap (e.g., Jaffee, 1989).

In Model 4, the effects of marital status, the presence of children, and occupation percent female are allowed to vary by the gender of the respondent. Thus, the significance tests for these three interaction terms assess whether the effects of these variables increase (or decrease) the odds of possessing workplace authority by a different amount for women and men (equivalently, these models test whether the net gender gap varies across levels of gender composition). Because of the coding of the gender variable (female = 1), the effect of percent female for men is simply the gender composition coefficient. For women, the effect is equal to the sum of the gender composition coefficient and the coefficient on the multiplicative term. Here, the interaction term is not statistically significant ($p > 0.10$). This suggests that both women's and men's odds of possessing workplace authority decrease at similar rates as female representation increases. This finding is at odds with those reported by Kraus and Yonay (2000), who report a greater female disadvantage in female-dominated occupations. Thus, Model 4 of Table II does not support the view that as female representation increases competition will be intensified, and women will face a greater disadvantage as a result (Blalock, 1967). In addition, the argument that the presence of many women in an occupation will help remedy women's disadvantage is not supported (Kanter, 1977). These results suggest that women's disadvantage largely results from the paucity of opportunities to exercise authority in female-dominated occupations—both men and women are similarly disadvantaged when employed in such occupations. However, because of occupational gender segregation, this translates into a large net disadvantage for women.

The inclusion of the interaction terms for Gender × Marital status and the Gender × Presence of children tests arguments based on the self-selection of women into positions with low levels of authority. Regarding the presence of children, these arguments posit that women with children self-select out of competition for jobs involving authority (Wright *et al.*, 1995:411). Thus, the effect of the presence of children on the likelihood of authority should be negative for women but close to zero for men. Likewise, predictions from human capital theory suggest that marriage may signal responsibility and stability to employers among men but not among women. This implies a marriage premium for men and a marriage penalty for women (Cohen, 2002). Model 4 contradicts both these views. There is no evidence that marriage or the presence of children alters the likelihood of possessing workplace authority differently for women and men.

The final model in Table II tests whether the negative effect of occupation gender composition depends on whether an occupation is national or local in scope. The results suggest that it does. For occupations in local labor markets, the effect of occupation percent female is statistically significant and negative (OR = .988). In national labor markets, however, the negative effect of occupation percent female is substantially steeper. Thus, everyone's odds of possessing work authority decrease with the representation of women. Importantly, this effect is magnified in the more desirable, national labor market occupations. Clearly, then, while the bias against female-dominated lines of work found in studies of earnings attainment also is found for workplace authority, treating all occupations

Table III. Determinants of Gender Gap in Workplace Authority, by Occupational Labor Market: Odds Ratios From Logistic Regression Models

Variable	Local Labor Market (Model 1)	National Labor Market (Model 2)
Main effects		
Gender (female = 1)	1.133	0.016**
Occupation percent female	0.992	0.926***
Gender interaction		
Gender × Occupation percent female	0.991	1.092***
Control variables		
Industry and human capital variables	Yes	Yes
Likelihood ratio chi-square[a]	101.9***	55.7***

[a]Test of improvement in model fit versus the intercept-only model.
*$p < 0.10$; **$p < 0.05$; ***$p < 0.01$ (two-tailed test).

as if they were part of one large labor market results in understating that penalty for those working in occupations that operate in the national labor market.

The Gender Authority Gap: Differences Across Labor Markets

Table III displays the results of regression models estimated separately for the national versus local labor markets. These models address whether the size of the gender gap varies by *both* occupation gender composition and occupational labor market. To simplify the presentation, only the coefficients for key independent variables are shown. In each model, the effect of gender composition for men is estimated by the main effect of occupation percent female. For women, the gender composition effect equals the sum of the main effect of percent female and the coefficient on the interaction term. Thus, comparing the gender-specific effects in each model indicates how the gender authority gap covaries with occupation gender composition.[9]

[9]Readers may wonder whether the gender gap is statistically significant at various points of the percent female distribution (particularly at the extremes). These tests are ill-advised because of the extremely small sample sizes at the extremes. For example, by definition, there are not enough men in female-dominated occupations, and not enough women in male-dominated occupations to perform a test with adequate power. Of course, one possible solution to this problem is to use a categorical, rather than continuous, measure of gender composition. For example, Kraus and Yonay (2000) distinguished between male-dominated (less than 27% female), gender-mixed (between 28 and 65% female), and female-dominated (more than 65% female) occupations. However, this solution is appropriate for larger data sets, especially where tests are not attempted separately for national and local labor markets. Thus, we base the analyses on the gender-specific slopes, rather than on significance tests with unacceptably large standard errors. This also avoids using arbitrary cut-offs for distinguishing between male-dominated, gender-mixed, and female-dominated occupations. Since the 1991 GSS is one of the only nationally representative samples that includes an acceptable authority measure,

The first model in Table III applies to local labor markets. The effect of gender composition for men is not statistically significant, meaning that men's likelihood of possessing work authority is unrelated to the occupation's gender composition. However, while the coefficient on the interaction term is also not statistically significant, the effect of gender composition for women—which equals the sum of the main effect and the interaction effect—*is* statistically significant ($p < 0.001$). This provides some evidence that in local labor market occupations the gender gap is larger in female-dominated occupations than in those that are male-dominated. This supports Kraus and Yonay's (2000) finding of a larger gender authority gap in female-dominated occupations.

However, a switch occurs in the national labor market (see Model 2 of Table III). Here, *men's* odds show a statistically significant association with occupation percent female, with men's odds of possessing authority declining as the percent female increases (OR = .926, $p < 0.01$). For women in the national labor market, however, there is not a statistically significant relationship between occupation percent female and the odds of possessing workplace authority. However, the negative main effect of gender is extremely large and statistically significant (OR = .016, $p < 0.05$). This means that in male-dominated, national labor market occupations, men enjoy a marked net advantage over women. Because the effect of percent female is statistically significant and negative for men, their advantage shrinks as percent female increases. Thus, in national labor markets, the gender authority gap is negatively related to occupation percent female.

DISCUSSION AND CONCLUSION

Although prior research has shown that female-dominated occupations are characterized by low levels of authority, we show how this relationship varies depending on whether the occupation is national or local. In doing so, we offer a new method for classifying occupational categories according to whether workers tend to compete for jobs in the national labor market or in a local labor market. Theoretically, this is important because it defines the arena in which competition occurs and improves the specificity of our models. This classification allows us to extend previous empirical work on the impact of occupational gender segregation on various dimensions of workplace gender inequality (e.g., England, 1992; Nelson and Bridges, 1999; Petersen and Morgan, 1995; Reid, 1998;

we think this weakness is minor when weighed against the advantages gained from using these data.

Tomaskovic-Devey, 1993), by showing how processes leading to inequality depend on the occupational labor market. Our central findings are summarized below.

First, as in the case of earnings, typically female occupations are disadvantaged in terms of opportunities for authority. To the extent that authority is one aspect of work used by employers to determine wage levels, this finding may be unsurprising. However, those interested in explaining the authority gap should be interested in the strength of the gender composition effect. Thus, with respect to work authority, the results are consistent with a status composition process (Tomaskovic-Devey, 1993) through which typically female occupations offer fewer chances to exercise authority because they are performed by women. However, it may be the case that men monopolize positions of authority, closing them off to women. Such a status closure processes would also produce a strong net effect of occupational gender composition (Tilly, 1998).

The findings also indicate that the penalty accruing to female-dominated occupations is not uniform across occupational labor markets—whether an occupation operates in a national labor market or in a more circumscribed, local labor market affects the degree to which female-dominated occupations are penalized in terms of opportunities for authority. Among national labor market occupations, the penalty against typically female-dominated occupations is significantly magnified. Additionally, because we analyze cross-sectional data, we cannot completely rule out the possibility that the causality is reversed—it might be that women are channeled into occupations with few opportunities to exercise authority, rather than occupations offering little authority because they are performed mostly by women.

The negative effect of occupation percent female could be due to differences in skill requirements found in different occupations rather than bias against female-dominated occupations. In other words, is there an authority penalty imposed on occupations where women come to dominate, or are women channeled into occupations that, from the start, require few skills and have few chances to exercise authority? In supplementary analyses (not reported), we added two occupational-level control variables taken from the *Dictionary of Occupational Titles* (U.S. Department of Labor, 1977) to the models to explore this possibility. Those contextual variables, common to studies of gender composition effects (England, 1992; Tam, 1997), tap the average educational requirements (GED) and amount of specific vocational training (SVP) required to perform the work requirements adequately. Their inclusion did not alter the findings substantially. This increases our confidence that the negative effect of percent female in an occupation on work authority is attributable to who is performing the work in a given occupation. If we interpret the negative effect of occupation percent female as one form of gender bias, we then can conclude that this type of bias is markedly stronger when there is more at stake—in the more desirable national labor market occupations. This suggests that occupation gender composition means something different depending on the occupational labor market; however, our models do not illuminate the specific mechanism for this.

Second, following the application of Blalock's (1967) competition hypothesis to occupational composition, we hypothesized that occupational-level changes in gender composition would primarily affect the gender gap in those occupations that operate in the national labor market. This prediction would have been supported by a more rapidly increasing gender gap in authority as occupation percent female increases for those in the national labor market, compared to occupations in local labor markets. This prediction was unsupported.[10] In national labor market occupations, we find a smaller gender gap in female-dominated occupations than in those that are male-dominated. Among those occupations, women's odds of possessing work authority are independent of gender composition, and men's odds are lower than in male-dominated occupations. We also find some evidence that the gender gap is positively related to female representation among local labor market occupations. Further research should address this unexpected finding, which suggests that varying levels of advantage or disadvantage accruing to "tokens" depend not only on one's gender, but also on the type of occupation in which one is employed.

To test whether disparities in the interaction effects in the models in Table III are due to differences in average occupational education, and not labor market scope, we fit those models with an occupational-level measure of education (the proportion of workers in the occupation with a B.A. degree or higher). The results in fact do not change appreciably with the addition of this variable, thereby increasing our confidence that our measure captures important aspects of labor market scope and not educational levels. As an additional check on whether the effect of educational level and labor market scope are conflated, we reran our models with only the highly

[10]Although this finding does not appear to support Blalock's (1967) competition hypothesis, this result might stem from the application of the competition hypothesis to occupational composition. Even if men's perceptions of threat are heightened by an increase in women's occupational representation, some might argue that they lack a specific mechanism for keeping women out of authority positions. In fact, Blalock (1956:585) states that "keeping subordinate groups out of clubs or neighborhoods may be easier than controlling wages or limiting suffrage." However, to the extent that men can exclude women from jobs that have authority, a significant net effect of occupation gender composition will result. Nonetheless, this does not explain the findings in Table III.

educated occupations. The main results of the analysis were not substantially different.

This study is not without weaknesses. First, our authority measure ignores variation in the meaning (and amount) of authority across the two types of occupational labor markets. For example, there may be important differences in the typical number of subordinates supervised, as well as the kind of authority exercised (for example, authority to hire or fire employees or make budgetary decisions). Our measure does not tap these key dimensions of work-based authority.

Second, tests of theories regarding work-based competition and tokenism may be better served with gender composition measures that describe jobs (specific work titles in a work establishment) rather than occupations (broad collections of jobs wherein workers perform similar work across employment contexts). Considering job-level gender composition could also be important because the penalty for working in a typically female *occupation* may be offset by working in a male-dominated *job*. Additionally, processes determining which jobs have authority and who gets matched to those jobs occur primarily at the more localized job level, rather than at the broad occupational level. Thus, one fruitful area for future research would be to compare the effects of job versus occupation gender composition. This research would do more to reveal the mechanisms producing unequal outcomes in work authority and also provide better tests of the theories addressed here.

Additionally, more work devoted to classifying the labor market scope of occupational categories would be beneficial. Although some occupations may remain difficult to classify, a multimethod approach to labor market scope would be positive step. For example, the existence of an occupation-specific association or organization could be used to indicate whether an occupation operates in the national labor market. Additionally, a random sample of employers could be asked whether they recruit locally or nationally for particular occupations. Carefully triangulating among these various measures would result in increased confidence about the relevant labor market for an occupation.

These findings have implications beyond the study of workplace authority. A wide range of research relating to gender and race-based inequality centers on relating occupation gender and race composition to various outcomes. Our findings suggest that ignoring variation in the relevant labor market for different occupations could mask important differences. This article should be viewed as a preliminary step in a larger research agenda focused on understanding not only how the contours of gender stratification are shaped by factors such as demographic composition, but also how the effect of such factors is contingent on other occupational attributes.

APPENDIX A

Five Most Male-Dominated and Female-Dominated Occupations, by Occupational Labor Market Type

	Percent female
Local Labor Markets	
Male-Dominated Occupations	
1. Bus, truck, and stationary engine mechanics	0.88
2. Heavy equipment mechanics	1.10
3. Brickmasons and stonemasons	1.25
4. Heating, air conditioning, and refrigeration mechanics	1.33
5. Electrical power installers and repairers	1.40
Female-Dominated Occupations	
1. Secretaries	98.69
2. Teachers, prekindergarten, and kindergarten	97.80
3. Receptionists	95.74
4. Typists	94.35
5. Registered nurses	94.31
National Labor Markets	
Male-Dominated Occupations	
1. Mechanical engineers	5.26
2. Civil engineers	6.98
3. Aerospace engineers	8.12
4. Electrical and electronic engineers	9.97
5. Podiatrists	11.27
Female-Dominated Occupations	
1. Speech therapists	91.14
2. Elementary school teachers	78.44
3. Health specialties teachers	75.91
4. Physical therapists	75.51
5. Therapists, n.e.c.	72.33

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