# INDIVIDUALS, JOBS, AND LABOR MARKETS: THE DEVALUATION OF WOMEN'S WORK

**PHILIP N. COHEN**
*University of California, Irvine*

**MATT L. HUFFMAN**
*University of California, Irvine*

*Although abundant evidence documents pay penalties for female-dominated jobs, there is also substantial variation in gender inequality across U.S. metropolitan areas. These lines of research are united by exploring whether occupational gender segregation at the labor market level exacerbates the wage penalty associated with female-dominated jobs, and investigating the association between gender composition and the size of within-job gender gaps. Results show that the penalty accruing to female-dominated jobs is weaker in more integrated labor markets, but only among men, and that labor market integration does not significantly influence the association between the gender composition of jobs and within-job inequality. Further, even women in completely segregated jobs benefit from a context of occupational integration. It is concluded that, although gender devaluation is widespread and systematic, variation in gender composition effects across local contexts is an important dimension of gender inequality.*

**WORK DONE** primarily by women is rewarded less than work done by men. This has been documented for broad occupational categories (England 1992; England et al. 1988) and for specific job titles in work establishments (Baron and Newman 1989; Bridges and Nelson 1989; Huffman and Velasco 1997; Petersen and Morgan 1995; Tomaskovic-Devey 1993b). Although economists (Killingsworth 1989) explain this result by pointing to differences in skills, working conditions, and supply and demand factors, sociologists assert that pay penalties result from widespread cultural devaluation of women's work (England 1992). Indeed, implicit in some sociological work on gender inequality is the notion that gender inequality is universal, reflecting broad biases against women's work (Acker 1990; England 1992). We have no reason to doubt this, especially in light of research showing that both women and men tend to assign more worth (Deaux 1985; McArthur 1985) and prestige (Bose and Rossi 1983) to work performed by men. Additionally, those skills closely associated with women's work, such as nurturance, are systematically underrewarded (England et al. 1994; Folbre 2002).

However, the existence of a general bias against female-dominated work does not preclude the possibility of significant variability in devaluation across various contexts, such as organizational settings (Baron and Newman 1989; Huffman and Velasco 1997; Tomaskovic-Devey 1993b). Although the tendency to devalue women's work may be systematic, the mechanisms producing it are often local (Tomaskovic-Devey 1995). We know that there is significant variation in gender inequality across local areas (Cotter et al. 1997; Lorence 1992). And Jacobs and Blair-Loy (2001) argue that "the more variation across metropolitan areas, the more importance theories must place on local factors" (p. 354). However, existing research has failed to investigate whether gender de-

Direct all correspondence to Philip N. Cohen, Department of Sociology, University of California, Irvine CA 92697 (cohenp@uci.edu). This research was supported by a grant from the University of California Institute for Labor and Employment. We thank Reeve Vanneman and his colleagues for sharing their Metropolitan Area data file with us, and Richard Congdon for his valuable computing suggestions. We also thank Frank D. Bean and Calvin Morrill for helpful comments on an early version of this paper. The authors contributed equally to this manuscript.

valuation is conditioned by the local labor market context. If this is the case, the pattern of this variability may shed light on how gender inequality is structured and thus reproduced.

Our analysis makes several unique contributions to the literature on gender composition and wages by uniting research on wage penalties for "women's work" with the literature on spatial variation in occupational segregation—the tendency for men and women to work in different occupations. Most significant, we test whether occupational gender segregation at the level of the local labor market exacerbates both the penalty for working in female-dominated jobs and the wage gap within jobs. Thus, we offer the first test of spatial variation in the wage effect of the gender composition of jobs, while paying special attention to the role of gender segregation at the local labor market level.[1] The results have broad implications for understanding the structure of gender inequality in U.S. labor markets.

## PREVIOUS RESEARCH

### GENDER COMPOSITION EFFECTS

Many scholars link the high level of occupational gender segregation (Baunach 2002; Tomaskovic-Devey, Kalleberg, and Marsden 1996; Wells 1999) to wage inequality (England 1992; Huffman and Velasco 1997; Petersen and Morgan 1995; Tomaskovic-Devey 1993b). The overrepresentation of men in better-paying jobs produces a strong association between gender composition and average pay levels. After adjustments are made for the differences in the skill and educational requirements between female- and male-dominated jobs, advocates of comparable-worth policy see the remaining association between gender composition and rewards as evidence of gender-bias in how wages are attached to work roles (England and Dunn 1988; Nelson and Bridges 1999).[2] However, the effect of gender composition on average rewards is distinct from inequality resulting from pay differences occurring *within* jobs. Thus, the overall gender pay gap results from two sources: the differential distribution of women and men across jobs and occupations that vary with respect to average pay, and within-job pay differences.

Specifically, England, Reid, and Kilbourne (1996) find a wage difference of 7 to 19 percent between male- and female-dominated occupation-industry cells. Kilbourne et al. (1994) and Macpherson and Hirsch (1995) report a difference of approximately 5 percent. Similarly, England et al. (1988) find wage differences between male- and female-dominated occupations ranging from 8 to 11 percent, depending on the model specification.[3]

Although data are more readily available for occupations than for jobs, those studies that do examine jobs find a much stronger effect of gender composition on wages.[4] This is at least partially explained by the fact that jobs are more segregated than occupations (Bielby and Baron 1986; Tomaskovic-Devey 1995). Baron and Newman (1990) report that the difference in starting wages between male- and female-dominated civil service jobs is approximately 30 percent, and Huffman, Velasco, and Bielby (1996) find the typical earnings in all-female jobs to be about

[1] In previous research, we found that gender composition effects on average wages are stronger when jobs are embedded in gender-segregated local labor markets (Cohen and Huffman 2003). Our present analysis, using 1990 Census microdata, includes a much larger sample of jobs and applies a more stringent test with controls for individual characteristics.

[2] Lower pay in female-dominated positions may also reflect women's crowding into a limited range of occupations (Bergmann 1974; Jacobsen 1994). In practice, crowding and devaluation are difficult to distinguish; we refer to gender composition and devaluation effects interchangeably.

[3] In contrast, Tam (1997) reports no significant negative effect of occupation gender composition on wages. For a counterargument, see England, Hermsen, and Cotter (2000).

[4] Job-level studies can test whether the wage effect of gender composition is conditioned by contextual factors such as organizational age, size, and the degree of formalization of personnel practices. For example, Baron and Newman (1990) find stiffer penalties associated with female representation among generic job titles, older jobs, and nonunionized jobs. However, Huffman and Velasco (1997) report this penalty to be relatively uniform across diverse establishments.

half of those in comparable all-male jobs. And perhaps most striking, both Petersen and Morgan (1995) and Tomaskovic-Devey (1993a) report that the gender composition of jobs accounts for over 90 percent of the gender wage gap.

Thus, studies conducted at both the occupational level and job level have shown that female representation negatively affects rewards, but the penalty accruing to female-dominated jobs is much larger than that associated with occupations. Because wage rates in the United States are attached to jobs, with no national system linking wages to broad occupations (Tomaskovic-Devey 1995:24), occupation-level wage penalties based on gender composition chiefly reflect the aggregation of job-level effects (Huffman, Velasco, and Bielby 1996; Tomaskovic-Devey 1995). However, the fact that job-level wage penalties are larger than occupation-level effects also results from local variation in wage-setting practices. Despite the importance of local factors in wage setting, they are not accounted for in national studies using occupations as the unit of analysis, or in job-level studies that do not assess the broader labor market context of jobs.[5]

If most of the wage gap is due to the gender composition effect, this implies that gender inequality is driven largely by the segregation of workers into jobs with different average pay levels. However, within-job inequality also contributes to the wage gap; it reflects processes that benefit particular groups within jobs.[6] Thus, while the gender composition of jobs may have a substantial effect on average pay levels (thereby contributing to between-job inequality), the within-job component of the wage gap may also be strongly correlated with jobs' demographic composition. However, theorizing about the direction of this association—whether larger gender gaps occur in male- or female-dominated jobs—has produced contradictory predictions.

On one hand, Kanter's (1977) theory of tokenism suggests that increased concentrations of women will *decrease* gender inequality. Kanter posits that men's perception of women is conditional on an organization's gender composition—when women's representation increases, women face less performance pressure, have more power than women in male-dominated work settings, and therefore suffer less discriminatory treatment (Blau 1977; Jacobs 1992; Kraus and Yonay 2000; Reskin, McBrier, and Kmec 1999).

On the other hand, the pattern of *more* inequality among female-dominated work settings is consistent with a "glass escalator" effect (Budig 2002; Williams 1992), whereby men in female-dominated occupations benefit from hidden structural advantages that enhance their career opportunities.[7] As Acker (1990) argues, "In contrast to the token woman, White men in women-dominated workplaces are likely to be positively evaluated and to be rapidly promoted to positions of greater authority" (p. 143). Kraus and Yonay (2000) report a more pronounced gender gap in workplace authority in female-dominated occupations, which they attribute to a variation of Blalock's (1967) "competition hypothesis"—that White racism is more severe in the presence of larger minority

[5] Another concern raised by many job- and occupation-level studies is that the negative effect of gender composition may be confounded with the fact that women are paid less than men across the board. Thus, what is attributed to the devaluation of jobs based on gender composition may in part reflect constant levels of discrimination against individual women. Studies performed at the job level that examine the association between average wages and percentage female (e.g., Huffman and Velasco 1997) are subject to this criticism because they cannot disentangle the between-job portion of the gender wage gap from inequality within jobs. Our research design permits us to assess both sources of the wage gap, so we can test whether the gender composition effect persists once the overall lower pay of women is accounted for.

[6] The 1963 Equal Pay Act was explicitly designed to address this type of wage discrimination (England 1992; Nelson and Bridges 1999).

[7] Williams's (1992) study is based on in-depth interviews of workers in only four predominantly female professions: nurses, school teachers, librarians, and social workers. Huffman et al. (1996) also find evidence supporting the view that men—but only those in managerial and professional occupations—are able to escape the earnings penalty associated with female representation.

populations. Thus, the largest gender gap should be found in female-dominated contexts, because in male-dominated occupations there is less competition over scarce resources, and therefore less gender discrimination.[8] In recent research on Japan, Aiba and Wharton (2001) find that women's wages are negatively affected by the percentage female in the job, whereas men's wages are largely unaffected. We also expect to find more within-job gender inequality in female-dominated jobs.

Although clearly widespread and systematic, we do not yet know the extent to which the dynamics that cause gender composition effects are produced locally. If the costs associated with high concentrations of women are based on the evaluation of positions—and the motivation or ability of men and women to compete for rewards—then it is likely that the outcome will depend in part on the context in which these actions occur. This variation may yield important insights into the mechanisms at work; as a result, it is important to address the role of local labor markets.

### *Reproducing Devaluation across Labor Markets*

Undoubtedly, a crucial step toward understanding gendered wage inequality was the observation made by Baron and his colleagues that there is significant variability in the devaluation of women's work that is linked to features of employing organizations (Baron, Mittman, and Newman 1991; Baron and Newman 1989, 1990). In a similar spirit, we ask: Does the tendency for female-dominated jobs to be penalized depend on characteristics of the labor market in which those jobs are embedded? In other words, does the devaluation of women's jobs vary spatially as a function of labor market characteristics?

Given that most wages in the United States are set at the job level in local labor markets, Tomaskovic-Devey (1995) notes that the relative influence of processes that favor or discourage gender devaluation should vary "from place to place and over time as a function of the local relations of production" (pp. 42–43). Jacobs and Blair-Loy (2001) specifically encourage researchers to adopt a multilevel approach to studying gender and race composition effects, so that contextual influences on inequality can be specified. At a higher level of abstraction, Britton (2000) convincingly argues that the assumption of a constant level of "gendering" of women's work roles is contradicted by research showing marked variation in this process and the associated penalties. Although Britton is mainly concerned with how organizational characteristics shape the gendering processes, she also states that "if one can identify those factors that are conducive to less gender segregation and inequality in organizational or occupational or *labor force* environments, then the possibility of replicating those conditions becomes much more realistic" (p. 423, italics added).

These specific appeals notwithstanding, prior research on the wage effects of gender composition has not examined how gender-based devaluation varies across local labor market contexts, despite a substantial body of literature suggesting the salience of such factors for other dimensions of gender inequality. For example, the importance of local context is found in research showing that levels of racial and gender inequality are sensitive to attributes of firms' local institutional environment, including the degree of support for equality of opportunity (Beggs 1995; Guthrie and Roth 1999). Other recent research on racial stratification has found significant variation in processes leading to inequality, such as income returns to education, across geographic regions of the United States (McCall 2000). Additionally, employers respond to changes in the legal environment with regard to gender inequality issues and workplace due process procedures (Edelman 1990; Kelly and Dobbin 1999). And, through local laws, cities can set more stringent employment discrimination standards than those provided by the federal government (Gold 1993; Gutman 1993).

Occupational gender segregation at the local labor market level has been shown to depress earnings for all women, including those in integrated occupations (Cotter et al.

[8] For example, Morgan (1998) reports that women, who make up only 8 percent of all engineers, earn 97 percent of male engineers' earnings.

1997). Cotter et al. argue (pp. 715–16) that in labor markets with more gender equality in occupational allocation, normative expectations, the balance of managerial power, and market pressures may result in increased earnings and better jobs for all women. Although the mechanisms for this remain unclear, their findings could account for some of the job-level gender-composition effect found in previous research.

## HYPOTHESES

Our hypotheses treat both between- and within-job inequality. Each applies after controlling for variables at the individual, job, and (where appropriate) labor market levels. Although our first two hypotheses have been addressed in previous research, our multilevel research design provides a more stringent test. The first hypothesis represents the basic gender composition effect:

*Hypothesis 1*: Average wages will be lower in jobs with high female representation.

Second, we hypothesize that female representation will also increase within-job gender inequality:

*Hypothesis 2*: Within-job inequality will be higher in jobs with a high proportion of women.

Hypothesis 2 is consistent with the "glass escalator" hypothesis, which posits that men in female-dominated jobs escape some, or all, of the gender composition penalty, resulting for them in higher pay and faster promotions.

After testing these basic composition effects, we move to our main questions. Do labor markets in which men and women are less segregated into different occupations exhibit weaker gender composition effects? Our third hypothesis concerns the *between-job* wage gap:

*Hypothesis 3*: The negative effect of proportion female in jobs on average wages will be weaker in integrated labor markets.

This would be the case if in integrated markets women had more bargaining power because of their greater range of options, if women more frequently were found in decision-making positions and on average judged women's work more fairly, or if women's visibility in a wide range of positions undermines local discriminatory attitudes about women's roles and the worth of female-dominated work (Cohen and Huffman 2003; Cotter et al. 1997).

Our fourth hypothesis examines the effect of labor market integration on *within-job* gender inequality. Formally,

*Hypothesis 4*: The positive relationship between the proportion female in jobs and the wage gap between men and women will be weaker in integrated labor markets.

This hypothesis suggests that the tendency for men to rise to the top of female-dominated jobs (the "glass escalator") is more pronounced in segregated labor markets, perhaps because women in those markets possess less authority and overall weaker bargaining positions. The opposite might be true, however, if the devaluation of women's work in segregated markets undermines men's "glass escalator." That could occur if the mechanism for the apparent glass escalator is actually a "status inconsistency" penalty for working in female-dominated jobs (Cassidy and Warren 1991)— and if such an effect is stronger in segregated labor markets where there is more gender devaluation. Previous research does not strongly suggest one of these alternatives over the other.

These hypotheses refer to the gender composition of jobs and the gender segregation of labor markets. We should note, however, that racial/ethnic-specific gender composition effects may occur, especially at the job level. That is, the proportion of Black or Latina women in a job—beyond the simple gender composition—may influence earnings (Catanzarite 2002; Reid 1998). Because attending to this possibility requires additional theorizing—and would further complicate the multilevel models—we defer this question to a later project.

## DATA, MEASURES, AND MODELS

### DATA

Our primary data source is the 1990 Census (5-percent Public Use Microdata Sample).

**Table 1. Descriptive Statistics for Variables Used in the Analysis: Men and Women, Ages 25 to 59, 1990**

| Level of Analysis/Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| INDIVIDUALS (N = 1,920,100) | | | | |
| Wage (in 1989 $U.S.) | 14.14 | 11.93 | 1 | 249.80 |
| Wage (ln) | 2.44 | 0.64 | 0 | 5.52 |
| Female | .49 | .50 | 0 | 1 |
| Own children | .92 | 1.14 | 0 | 18 |
| Foreign-born | .13 | .34 | 0 | 1 |
| Disabled | .03 | .17 | 0 | 1 |
| *Race/Ethnicity:* | | | | |
| Latino | .09 | .29 | 0 | 1 |
| Asian | .04 | .19 | 0 | 1 |
| Black | .10 | .30 | 0 | 1 |
| Other | .01 | .07 | 0 | 1 |
| Education | 13.76 | 2.93 | 0 | 20 |
| Potential experience | 19.26 | 9.84 | –1 | 53 |
| (Potential experience)$^2$ | 467.63 | 434.53 | 0 | 2,809 |
| Married | .67 | .47 | 0 | 1 |
| Weekly hours (ln) | 3.67 | .33 | 0 | 4.60 |
| JOBS (N = 62,322) | | | | |
| Proportion female | .47 | .34 | 0 | 1 |
| Proportion foreign-born | .10 | .15 | 0 | 1 |
| *Race/Ethnicity:* | | | | |
| Proportion Black | .09 | .13 | 0 | 1 |
| Proportion Latino | .08 | .15 | 0 | 1 |
| Proportion Asian | .03 | .08 | 0 | 1 |
| Proportion other | .01 | .02 | 0 | .49 |
| *Requirements:* | | | | |
| General educational development | 3.77 | .82 | 1.556 | 6 |
| Physical demands | 1.66 | .81 | 0 | 3.93 |
| Standard vocational preparation | 5.35 | 1.50 | 1.710 | 8.51 |
| *Industry:* | | | | |
| Agriculture | .01 | .12 | 0 | 1 |
| Mining | .01 | .07 | 0 | 1 |
| Construction | .06 | .23 | 0 | 1 |
| Manufacturing | .19 | .39 | 0 | 1 |
| Transportation | .07 | .26 | 0 | 1 |
| Wholesale trade | .05 | .21 | 0 | 1 |
| Retail trade | .16 | .37 | 0 | 1 |
| Business services | .05 | .21 | 0 | 1 |
| Personal services | .03 | .17 | 0 | 1 |
| Entertainment | .02 | .12 | 0 | 1 |
| Professional services | .22 | .41 | 0 | 1 |
| Public services | .06 | .24 | 0 | 1 |
| Finance, insurance and real estate | .08 | .26 | 0 | 1 |

*(Continued on next page)*

*(Table 1 continued from previous page)*

| Level of Analysis/Variable | Mean | Standard Deviation | Minimum | Maximum |
|---|---|---|---|---|
| METROPOLITAN AREAS (N = 261) | | | | |
| Population (ln) | 12.71 | 1.08 | 10.95 | 16.78 |
| Proportion unemployed | .06 | .02 | .03 | .14 |
| Proportion Black | .11 | .11 | .00 | .46 |
| Proportion Hispanic | .07 | .14 | .00 | .94 |
| Proportion durable goods manufacturing | .10 | .06 | .01 | .32 |
| Female labor demand | .46 | .02 | .40 | .52 |
| *Region:* | | | | |
| Northeast | .13 | .34 | 0 | 1 |
| Midwest | .25 | .43 | 0 | 1 |
| South | .46 | .50 | 0 | 1 |
| West | .17 | .37 | 0 | 1 |
| Net internal migration | .00 | .05 | –.15 | .26 |
| Gender integration | .52 | .16 | 0 | 1 |

At the individual level, we include civilian workers in the prime working ages of 25 to 59, who are less likely to be students or retired (Fossett and Cohn 1995). We also limit the sample to those who earned between $1 and $250 per hour in 1989 and who were not self-employed (McCall 2001b). This selection yields the individuals at the first level of our analysis.

The second level of data is the job. Both Budig (2002) and England et al. (1996) use detailed occupation-industry cells as "jobs," but their job proxy does not include a geographic component. On the other hand, Hirsch and Schumacher (1992) construct cells from the less detailed, two-digit occupation and industry categories in conjunction with the four major census regions, which they consider a proxy for labor markets. The latter approach offers less occupation-industry detail, but does include a geographic component, albeit a rough one.

We combine these two approaches by including the detailed occupation and industry categories from the census with a more detailed local labor market dimension, making the cells a closer approximation of local jobs. To construct jobs, we assign each respondent to an occupation-industry-metropolitan area cell (for example, health technicians in hospitals in Las Vegas), using the three-digit occupation and industry codes, and 261 metropolitan areas from the file constructed by Cotter et al. (1997).[9] Although clearly not a perfect proxy for jobs such as may be observed in individual workplaces, our simulated jobs approximate such a measure while taking advantage of the larger sample size afforded by decennial census data.

To create the job-level data file, we used the individual data before imposing the age restriction, allowing workers outside that age range to contribute to the aggregate characteristics of jobs (such as proportion female). We excluded those jobs that had fewer than 10 incumbents before the age re-

[9] Cotter et al. (1997:716–17) constructed a file from the 1993 U.S. Census definitions of metropolitan areas, which uses metropolitan areas and Consolidated Metropolitan Areas (e.g., Washington-Baltimore) when commuting patterns indicate labor market integration over those areas. In New England, the file uses New England County Metropolitan Areas, which are more compatible with metropolitan areas in the rest of the country. Six small metropolitan areas were combined with nearby metropolitan areas, and one (Jacksonville, NC) was excluded because it was dominated by a military installation. The resulting file includes 261 metropolitan areas, representing approximately four-fifths of the U.S. population.

**Table 2. Job Characteristics at the Bottom, Median, and Top of the Wage Distribution: Men and Women, Ages 25 to 59, 1990**

| Wage Distribution/ Metropolitan Area | Occupation | Industry | Median Wage | Proportion Female | Median Wage Male | Median Wage Female | Female/ Male Ratio | Cell Size |
|---|---|---|---|---|---|---|---|---|
| *Low* | | | | | | | | |
| Laredo, TX | Household servant | Private households | 2.40 | .95 | 4.55 | 2.40 | .53 | 22 |
| El Paso, TX | Household servant | Private households | 2.77 | 1.00 | — | 2.77 | — | 64 |
| Brownsville, TX | Household servant | Private households | 3.00 | 1.00 | — | 3.00 | — | 55 |
| San Antonio, TX | Maid/ houseman | Building services | 3.00 | 1.00 | — | 3.00 | — | 15 |
| Brownsville, TX | Hairdresser | Beauty shops | 3.11 | 1.00 | — | 3.11 | — | 26 |
| *Median* | | | | | | | | |
| Las Vegas | Health technician | Hospitals | 11.10 | .68 | 10.58 | 11.58 | 1.10 | 19 |
| Philadelphia | Licensed practical nurse | Nursing facilities | 11.10 | .96 | 12.62 | 11.05 | .88 | 92 |
| Rochester, NY | Furniture sales | Furniture stores | 11.10 | .58 | 12.26 | 9.66 | .79 | 19 |
| Los Angeles | Laborer | Aircraft and parts | 11.10 | .10 | 11.14 | 10.38 | .93 | 29 |
| Canton, OH | Registered nurse | Nursing facilities | 11.10 | 1.00 | — | 11.10 | — | 18 |
| *High* | | | | | | | | |
| Milwaukee | Physician | Physician's offices | 51.99 | .00 | 51.99 | — | — | 17 |
| Jacksonville, FL | Physician | Physician's offices | 52.14 | .13 | 52.14 | 22.06 | .42 | 15 |
| Providence | Physician | Physician's offices | 52.35 | .07 | 54.36 | 23.08 | .42 | 14 |
| Memphis | Physician | Physician's offices | 57.24 | .21 | 61.64 | 20.76 | .34 | 14 |
| New York | Manager | Oil and extraction | 68.35 | .07 | 54.68 | 68.49 | 1.25 | 15 |

*Note*: Metropolitan area and occupation and industry titles are abbreviated. Wages are reported in 1989 U.S. dollars.

striction was imposed. The final data set includes 1,920,100 individuals nested within 62,322 jobs. These jobs, in turn, are nested within 261 U.S. metropolitan areas. Descriptive statistics for each level of the analysis appear in Table 1.

To illustrate the specific nature of our job measure, Table 2 reports statistics for five jobs each at the top, median, and bottom of the wage distribution for jobs with at least 10 incumbents. This shows, for example, that the lowest-paid jobs in our sample are 95 to 100 percent female and pay around $3 per hour (based on self-reports of annual weeks worked and hours usually worked per week). At the median, the jobs are between 10 and 100 percent female and pay $11.10 per hour. At the top, the jobs are 0 to 21 percent female, paying more than $50 per hour. Overall, men work in jobs that average 26.4

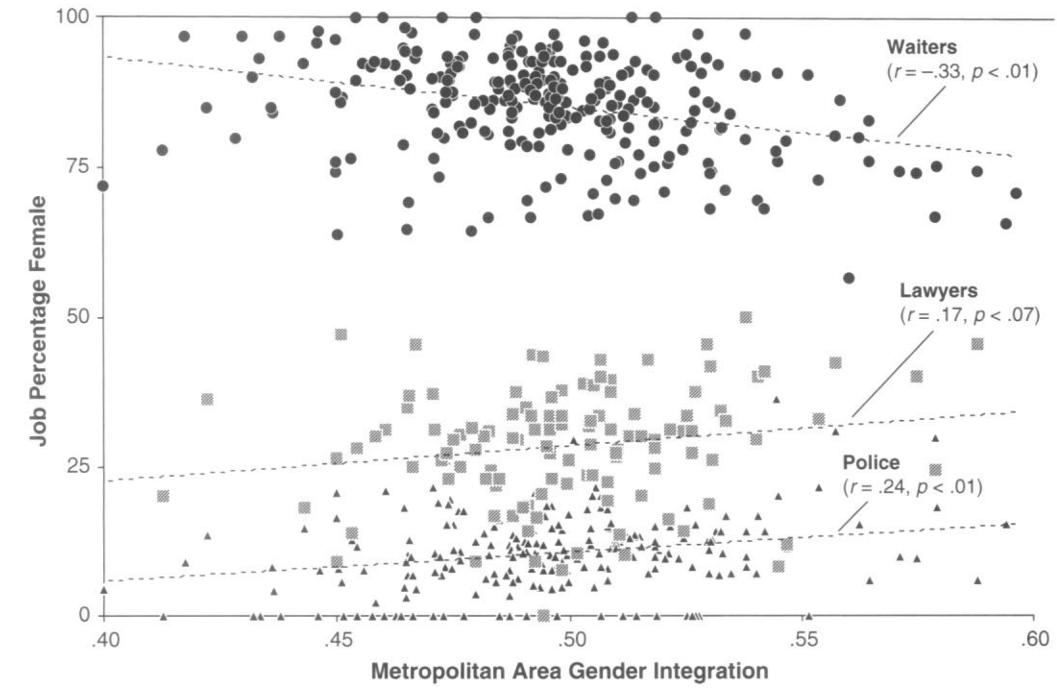


Figure 1. Gender Composition of Jobs by Metropolitan Area Gender Integration: Adult Men and Women, Ages 25 to 59, 1990

*Note*: Each job is from one industry: waiters (eating and drinking places), lawyers (legal services), police and detectives (justice, public order and safety); two-tailed tests.

percent female, and women are in jobs that average 70.9 percent female.

The geographic component of our job measure is crucial and unique to this study. We note that the gender composition of national occupation-industry cells (see Budig 2002; England et al. 1996) varies substantially across labor markets, partly as a function of occupational segregation at the labor market level. Because some of the variation in jobs' gender composition is related to local conditions, models using national occupation-industry cells may be underspecified. For many jobs, especially those that are male- or female-dominated, local gender composition is correlated with occupational integration at the labor market level. That is, women in occupation-industry cells that are segregated at the *national* level may be in less segregated jobs if their *local* markets are more integrated. To illustrate this, Figure 1 plots the gender composition of three industry-occupation cells by the level of occupational gender integration across labor markets. For waiters/waitresses in eating and drinking places, lawyers in legal services, and police and detectives in the justice, public order, and safety industry, the gender composition of the local job is closer to balanced in metropolitan areas that are more integrated overall.[10] Although not surprising, this fact underscores the need for considering important labor market characteristics in a three-level model; otherwise these effects remain undetected in analyses of wage variation.

[10] We selected these jobs to represent a range of gender compositions and wage levels. Note that this relationship does not hold for all jobs. Doctors in hospitals ($r = -.05$) and secondary school teachers ($r = -.09$), for example, show no significant correlation between gender composition at the job level and labor market occupational integration. One reason that some occupations do not show this correlation might be because the markets for those occupations in fact are national. Identifying the extent to which the labor markets for specific occupations are national versus local requires additional research (Huffman and Cohen 2003).

### *MEASURES*

At the individual level, the primary variable of interest is a binary indicator of respondent's *gender* (female = 1). We also control for a set of race/ethnicity dummy variables (1 = yes) indicating respondents who are *Black*, *Latino*, *Asian*, or *other*; and whether the respondent is *married*, *disabled*, or *foreign-born* (1 = yes). Continuous individual-level variables include *education* (in years of school completed), *potential labor market experience* (age minus years of education minus 6) and its *square*, the number of *householder's own children* present in the home, and the natural logarithm of *weekly hours worked* for each respondent.

Most of the variables at the job level are derived from data for the incumbents in each job. These include the main independent variable at the job level—*job proportion female*. We control for the proportions *Black*, *Latino*, *Asian*, and *other*, as well as proportion *foreign-born*. Following prior research on gender composition effects (England 1992; England et al. 1994; England et al. 1996), we also control for several variables appended to jobs from other sources. These include occupational characteristics from the *Dictionary of Occupational Titles* data set. Specifically, we use standard vocational preparation, general educational development, and the physical demands scale. Standard vocational preparation taps the amount of training time needed to learn the techniques, and obtain the information necessary for average performance on the job (high values represent a longer period of time required to acquire these skills). As such, this variable can be thought of as a measure of occupation-specific human capital (Tam 1997). In contrast, general educational development measures "the typical requirement of the occupation for schooling that is not vocationally specific" (England et al. 2000:1742). High values denote those occupations requiring relatively high levels of general educational development. The value of the physical demands scale is the average computed across five individual physical demands factors: stooping, climbing, reaching, talking, and seeing. High values of this variable indicate increased levels of physical demands. We also include dummy variable controls for each of the 13 broad industrial categories to capture systematic pay differences across industries (England et al. 1996).

At the metropolitan area–level, the primary independent variable is *occupational gender integration*. We use an adjusted dissimilarity index, taken from Cotter et al. (1997:730), rather than using the standard dissimilarity index (see Duncan and Duncan 1955). The adjusted measure yields the proportion of women who would have to change occupations so that the observed number of women was no larger or smaller than chance would predict. The interpretation of this measure is the same as the unadjusted dissimilarity index: the proportion of women or men that would have to change occupations in order to equalize the distribution of women and men across occupational categories. However, it uses a random distribution as the standard rather than absolute equality. The coding is reversed, so that high values of this variable indicate greater gender integration, rather than segregation. All labor markets in the United States exhibit marked occupational segregation. The most segregated metropolitan area is Houma, Louisiana (.390); the most integrated is Columbia, Missouri (.596). To facilitate interpretation of the results, we rescaled this variable, giving a value of 0 to the least integrated metropolitan area and a value of 1 to the most integrated.

Because of its association with a variety of outcomes related to gender inequality (Chafetz 1984; Cotter et al. 1998), we control for the *demand for female labor* implied by the occupational structure. This variable measures the share of the labor force that would be female if local occupations had the same percentage female as found in the country as a whole. It measures the degree to which local labor markets are slanted toward female-dominated occupations. A set of metropolitan area-level controls includes variables reflecting basic economic structure and historical conditions: the proportion of the labor force employed in *durable goods manufacturing* and dummy variables indicating each of the four census *regions*. Other variables reflect local economic conditions: the net percentage change in the population resulting from 1985–1990 *internal migration* (a proxy for long-term regional economic vitality), and the *unemployment rate* (for short-term vitality). Finally, to measure the local demographic structure we use the *size of the labor force* (logged), the proportion of the population that is *Latino*, and the proportion *Black*. We refer to individual-level variables as *Level-1* variables, job-level variables as *Level-2* variables, and labor-market variables as *Level-3* variables.

### STATISTICAL MODELS

Because our substantive questions hinge upon interactions between levels of analysis, which necessitates the use of nested data (individuals nested within jobs nested within metropolitan areas), we estimate a series of three-level hierarchical linear models (see Bryk and Raudenbush 1992; Snijders and Bosker 1999).[11] Hierarchical models allow tests of variability in the regression coefficients across levels of analysis (Kanaiaupuni and Donato 1999), thereby allowing us to address whether the effect of job gender composition on wages is conditional on characteristics of the metropolitan area in which they are embedded. To date, these questions have not been asked in the literature on gender composition and wages.

Multilevel models allow variation in the outcome variable to be decomposed across the levels of analysis; thus, we can differentiate between wage variability among individuals within jobs, wage variability among jobs within metropolitan areas, and wage variability among metropolitan areas. (For details on decomposing the variance across the three levels of data, see Appendix A.) These models avoid violating the assumption of statistical independence among the error terms that results from nested data structures. In our data set, for example, individuals in the same job have identical values on all variables measured at the job and labor market levels. Similarly, jobs in the same metropolitan area share all metropolitan area-level characteristics. Hierarchical models provide accurate standard errors and significance tests when data are multileveled (Bryk and Raudenbush 1992; Guo and Zhao 2000).

We model wage variation at each level as a function of individual, job, and labor market characteristics. At the individual level (Level 1), our model is

$$Y_{ijk} = \pi_{0jk} + \pi_{1jk}(Female_{ijk}) + \pi_{2jk}a_{1ijk} + \ldots + \pi_{mjk}a_{mijk} + e_{ijk}, \qquad (1)$$

where $Y_{ijk}$ is the log wage of person $i$ in job $j$ in labor market $k$, and $\pi_{0jk}$ is the intercept for job $j$ in labor market $k$. Next, $\pi_{1jk}$ is the individual-level effect of gender, $a_{mijk}$ denote the $M$ individual-level control variables, and $\pi_{2jk}$ through $\pi_{mjk}$ are the associated individual-level regression coefficients. The control variables are all centered around their grand means. Finally, $e_{ijk}$ is the Level-1 random effect. The coefficient $\pi_{1jk}$ is of particular interest because it represents the net within-job wage gap. With $\pi_{1jk}$ in the model, $\pi_{0jk}$ represents the average wage for men at the mean of the controls.

Each Level-1 coefficient relating individual characteristics to wages can be modeled as either a random or fixed effect across jobs. In our models, only the Level-1 intercept and the coefficient for the female dummy variable are permitted to vary across jobs. Thus, our Level-2 model is

$$\pi_{0jk} = \beta_{00k} + \beta_{01k}(Proportion\ Female_{jk}) + \beta_{02k}X_{1jk} + \ldots + \beta_{0qk}X_{qjk} + r_{0jk}, \text{ and} \qquad (2a)$$

$$\pi_{1jk} = \beta_{10k} + \beta_{11k}(Proportion\ Female_{jk}) + \beta_{12k}X_{1jk} + \ldots + \beta_{1qk}X_{qjk} + r_{1jk}, \qquad (2b)$$

where $\beta_{00k}$ is the intercept for the job-level model in labor market $k$. In turn, $\beta_{01k}$ is the effect of job proportion female on $\pi_{0jk}$. Likewise, $\beta_{10k}$ is the job-level intercept for the effect of being female, $\pi_{1jk}$, and $\beta_{11k}$ is the effect of job proportion female on the Level-1 effect of being female (this is a cross-level interaction effect). If $\beta_{01k}$ is negative, Hypothesis 1 is supported for men—that is, men earn less in jobs with high proportions female. If the sum of $\beta_{01k}$ and $\beta_{11k}$ is negative, women earn less as female representation increases and Hypthothesis 1 is supported for women. If $\beta_{11k}$ is negative (assuming wages are lower for women), the gender gap is

[11] We estimate our models using the HLM program (Bryk and Raudenbush 1992), version 5.04. Estimates are derived through maximum-likelihood estimation.

larger in female-dominated jobs and Hypothesis 2 is supported. Finally, $X_{1jk}$ through $X_{qjk}$ denote the $Q$ control variables in each of the job-level models (which are centered around their grand means), and $\beta_{12k}$ through $\beta_{1qk}$ are the regression coefficients associated with these control variables. The Level-2 error terms are denoted by $r_{0jk}$ and $r_{1jk.}$

Each Level-2 coefficient relating job characteristics to Level-1 effects on wages can be modeled as either a random or fixed effect across labor markets. In our models, only the Level-2 intercept and the job proportion female coefficient are permitted to vary across labor markets. Thus, our Level-3 model is

$$\beta_{00k} = \gamma_{000} + \gamma_{001}(Occupational\ Integration_k) + \gamma_{002}W_{1k} + \ldots + \gamma_{00s}W_{sk} + u_{00k}, \quad (3a)$$

$$\beta_{01k} = \gamma_{010} + \gamma_{011}(Occupational\ Integration_k) + \gamma_{012}W_{1k} + \ldots + \gamma_{01s}W_{sk} + u_{01k}, \quad (3b)$$

$$\beta_{10k} = \gamma_{100} + \gamma_{101}(Occupational\ Integration_k) + \gamma_{102}W_{1k} + \ldots + \gamma_{10s}W_{sk} + u_{10k}, \text{ and} \quad (3c)$$

$$\beta_{11k} = \gamma_{110} + \gamma_{111}(Occupational\ Integration_k) + \gamma_{112}W_{1k} + \ldots + \gamma_{11s}W_{sk} + u_{11k}, \quad (3d)$$

where $\gamma_{000}$, $\gamma_{010}$, $\gamma_{100}$, and $\gamma_{110}$ are the Level-3 intercepts in models of the Level-2 coefficients; $\gamma_{001}$, $\gamma_{011}$, $\gamma_{101}$, and $\gamma_{111}$ are the effects of occupational gender integration on the Level-2 coefficients. If $\gamma_{011}$ is positive (and assuming Hypothesis 1 is true for men), gender integration attenuates the negative effect of job proportion female on the intercept, and Hypothesis 3 is supported for men. Subsequently, the test of Hypothesis 3 for women is provided by the sum of $\gamma_{011}$ and $\gamma_{111}$. Assuming job proportion female increases within-job inequality, Hypothesis 4—that occupational integration reduces the effect of gender composition on within-job inequality—is supported if $\gamma_{111}$ is positive. Coefficients for the $S$ Level-3 control variables ($W$) are denoted by $\gamma$, and are centered around their grand means. The Level-3 error terms are given by $u$ for each labor market, $k$.

## RESULTS

### *Job Composition Effects*

Coefficients from our multivariate analysis appear in Table 3.[12] The model intercept, with variable centering, represents hourly wages of men at the mean of all control variables and (in the later models) in an all-male job in the least integrated labor market. The intercept for the female coefficient in the lower panel shows the difference between male and female wages. The intercepts for the job proportion female variable show the net effect of a change from no women to all women in the job on wages for men and on the gender difference in wages. Level-3 coefficients show gender integration effects on the intercepts and job composition coefficients.

Model 1 shows the unadjusted difference between male wages and female wages, controlling only for the variation across jobs and labor markets. The coefficients show that men's average wage is \$10.33 ($e^{2.335} = 10.33$), while women's average is significantly lower, \$7.87 ($e^{2.335-.272} = 7.87$). Model 2 gives these predictions at the mean of the Level-1 control variables; here, men's wages are slightly higher than in the first model (\$10.70) and women's wages are similar (\$7.94). This result is conceptually similar to that obtained by estimating a Level-1 model that includes the individual-level controls and dummy variables to represent the 62,322 jobs and the 261 metropolitan areas.

Model 3 adds job-level variables, for tests of our first two hypotheses. In this model, the effect of job proportion female on the model intercept ($-.141$, $p < .001$) represents the change in men's earnings associated with an increase from 0 to 1 in the job proportion female. The effect for women is modified by the proportion female effect on the female dummy variable ($-.074$, $p < .001$). Thus, the effect of job gender composition on the model intercept can be interpreted as the between-job gender composition effect for men, and the sum of the two

[12] Because we hypothesize directional effects, we present one-tailed significance tests. Complete results from all models are available from the authors on request.

**Table 3. Hierarchical Linear Regression Coefficients for Hourly Wages: Men and Women, Ages 25 to 59, 1990**

| Variable | Model 1 | Model 2 | Model 3 | Model 4 |
|---|---|---|---|---|
| *Intercept* ($\pi_0$) | | | | |
| Intercept ($\beta_{00}$) | 2.335*** (.008) | 2.370*** (.008) | 2.436*** (.008) | 2.451** (.047) |
| Gender integration ($\gamma_{001}$) | — | — | — | –.034 (.091) |
| Job proportion female ($\beta_{01}$): | | | | |
| Intercept ($\gamma_{010}$) | — | — | –.141*** (.008) | –.207*** (.035) |
| Gender integration ($\gamma_{011}$) | — | — | — | .121* (.064) |
| *Female* ($\pi_1$) | | | | |
| Intercept ($\gamma_{100}$) | –.272*** (.003) | –.298*** (.003) | –.205*** (.004) | –.283*** (.023) |
| Gender integration ($\gamma_{101}$) | — | — | — | .157*** (.045) |
| Job proportion female ($\beta_{11}$): | | | | |
| Intercept ($\gamma_{110}$) | — | — | –.074*** (.007) | –.021 (.036) |
| Gender integration ($\gamma_{111}$) | — | — | — | –.105 (.070) |
| Control variables | None | Level 1[a] | Levels 1 and 2[b] | Levels 1, 2, and 3[c] |

*Note*: Numbers in parentheses are standard errors.

[a] Controls include own children, foreign-born, disabled, race/ethnicity, education, potential experience, married, and weekly hours.

[b] Controls include racial/ethnic proportions, proportion foreign-born, general educational development, physical demands, standard vocational preparation, and industry.

[c] Controls include population size, unemployment, proportion Black and proportion Hispanic, durable goods manufacturing, female labor demand, region, and net internal migration.

*$p < .05$ **$p < .01$ ***$p < .001$ (one-tailed tests)

coefficients is the between-job effect for women. The effect on the coefficient for the female dummy variable alone yields the within-job gender composition effect. (Note that in this model, although the labor market variables are not included, job-level effects are tested controlling for the variance across labor markets.)

Consistent with Hypothesis 1, the coefficients from Model 3 show that female-dominated jobs pay significantly less on average for both men and women, controlling for individual-level and job-level characteristics. Additionally, female-dominated jobs have significantly more within-job gender inequality than do male-dominated jobs, as predicted by Hypothesis 2.[13] In other words, men escape some of the gender composition penalty. In dollar terms, these effects are substantial: Average wages for men in the model fall from \$11.43 to \$9.92 as job proportion female rises from 0 to 1, a drop of

[13] Model 3 also finds (not shown) that jobs with higher representation of Blacks and Latinos pay less on average—net of controls at the individual and job level—but the gender gaps within those jobs are smaller. The *Dictionary of Occupational Titles* variables show that jobs in occupations requiring more education and vocational preparation pay higher wages, while those making more physical demands pay less. On the other

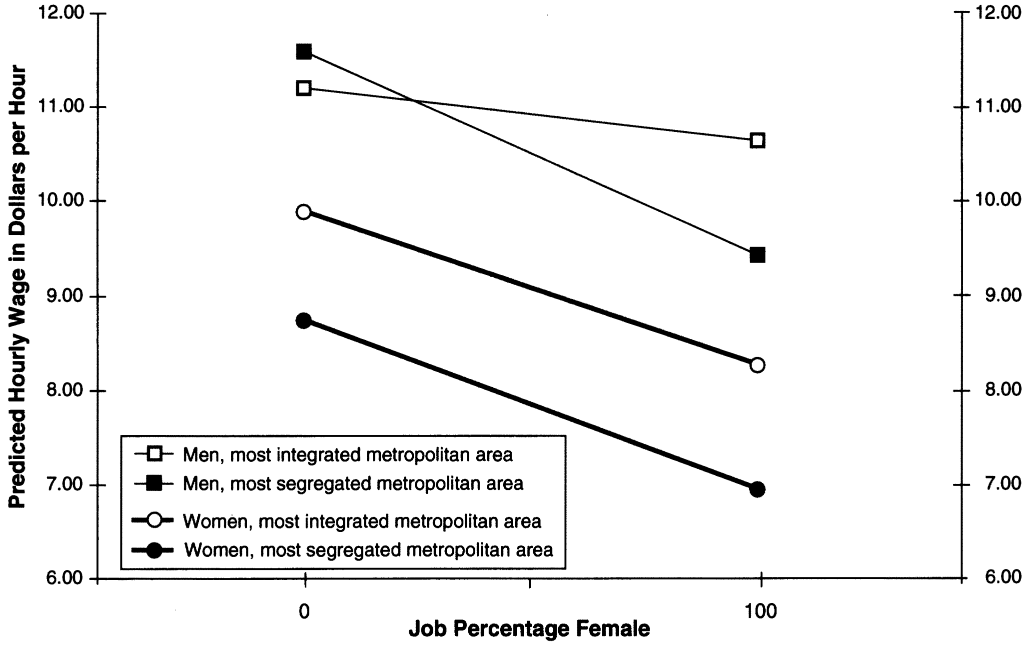


**Figure 2. Effect of Job Percentage Female on Hourly Wage, by Gender and Labor Market Occupational Segregation: Adult Men and Women, Ages 25 to 59, 1990**

13.2 percent. At the same time, women's wages are predicted to fall 19.3 percent, from \$9.31 to \$7.51. Within-job inequality increases as well, therefore, as women's predicted wages fall from 81.5 percent of men's wages in all-male jobs to 75.7 percent of men's wages in all-female jobs.[14] Model 3 also allows us to calculate the size of the gender gap if men and women both worked in perfectly integrated jobs (46 percent female), rather than at the observed level of segregation. In that case, the gender gap decreases by 26.7 percent.

### *LABOR MARKET INTEGRATION EFFECTS*

The final model in Table 3 tests our last two hypotheses. For men, gender integration at the labor market level reduces the between-job devaluation effect by more than half. In the most segregated labor market (at the intercept), the wage penalty for moving from an all-male job to an all-female job is –.207 ($p < .001$), but in the most integrated market that effect is reduced by .121 ($p < .05$). Thus, consistent with Hypothesis 3, between-job gender devaluation for men is significantly reduced in labor markets that are more gender integrated. However, although gender integration reduces the job composition penalty for men, the net effect of gender integration on the job composition effect for women is close to zero (.121 – .105 = .016).

Thus, with regard to Hypothesis 4, labor market gender integration does not significantly reduce the extent to which job proportion female increases within-job gender inequality. To the contrary, men pay a smaller penalty for being in female-dominated jobs when they are in integrated labor markets, but women do not. (However, we should note that results from a model without metropolitan area-level control variables are consistent with Hypothesis 4.)

The final model confirms, as Cotter et al. (1997) found, that in more integrated labor markets the net gender gap is significantly reduced. In the most segregated labor markets, the net gender difference is –.283, but

hand, jobs that require more education and that make more physical demands are associated with smaller within-job gender gaps, while high vocational preparation jobs have larger within-job gender gaps.

[14] Although women cannot be in all-male jobs, or men in all-female jobs, our data do include 99-percent male jobs and 99-percent female jobs.

in the most integrated labor markets, the net gender effect is –.283 + .157 = –.126. To simplify the presentation of the additive effects, we present predicted wages from the final model, converted to dollars per hour, in Figure 2. The figure shows predicted wages for men and women in male- and female-dominated jobs in the most and least integrated labor markets, at the mean of all control variables. Women in female-dominated jobs in the most segregated markets earn the lowest predicted wages ($6.95), while men in male-dominated jobs in the most segregated markets earn the most ($11.60).

Two patterns emerge from Figure 2. First, men in more integrated labor markets are conspicuous for the small gender composition penalty they pay. For both men and women, the proportion female effect is greater in segregated markets, but the difference for men is much more pronounced. If it is true that men in female-dominated jobs earn less because their status is compromised by the gender of their coworkers, this mechanism is much weaker in integrated labor markets.

Second, even women in the most segregated *jobs* are better off if they are in integrated *labor markets.* So although there are consistent penalties for women in female-dominated jobs, the context of labor market integration is still beneficial to all women (Cotter et al. 1997). Both of these findings underscore the fact that job-level gender composition tells only part of the gender-gap story. Labor market effects also are crucial.

## CONCLUSION

In this study, we have engaged unexplored questions about the interaction of labor market gender segregation and job-level gender composition effects, providing the first individual-, job-, and labor market level models of wage inequality. We bring together work on gender devaluation with the long tradition of research on patterns of inequality across local labor markets (e.g., Blalock 1967; Burr, Galle, and Fossett 1991; Cohen 1998; Semyonov, Hoyt, and Scott 1984). Specifically, we have asked whether occupational gender segregation may exacerbate the tendency for women's work to be devalued, and for gender gaps to be larger in female-dominated work settings. We draw several substantive conclusions from our results.

First, average wages are lower in jobs with high female representation. Given the many studies supporting this finding (e.g., Baron and Newman 1989; England 1992; Tomaskovic-Devey 1993b), this result may sound prosaic. However, we provide a more stringent test of the devaluation effect than previously reported because our models account not only for individual and job-level characteristics but also for variation across labor markets. Second, we find that there is greater within-job gender inequality in high proportion female jobs. Women in female-dominated jobs thus pay two penalties: Not only is the average wage in their jobs lower than that for comparable male-dominated jobs, they also earn less relative to men in the same jobs. These conclusions are consistent with our first two hypotheses.

The conclusions regarding Hypotheses 3 and 4 are less clear. The tendency for jobs to pay less as proportion female rises is less pronounced in gender-integrated labor markets, but only for men. We have no a priori explanation for why gender composition effects for men—but not for women—should be sensitive to the level of labor market integration. If Acker (1990) is correct that men in female-dominated workplaces are singled out for advancement, perhaps their "glass escalator" is undermined in segregated markets by a worse overall evaluation of female-dominated jobs. This could be interpreted as a "status inconsistency" penalty for men working in female-dominated jobs (Cassidy and Warren 1991), but only in the segregated labor markets where there is a more rigid devaluation of women's work.

On the other hand, we must be clear that women do benefit from gender integration at the labor market level. Although McCall (2001a) asserts that occupational segregation is not an important determinant of the gender wage gap, Cotter et al. (1997) find that all women, even those in segregated occupations, benefit from working in integrated labor markets. Our results confirm this finding, but also show an additional benefit from labor market integration: Even women in completely segregated *jobs* gain from working in labor markets with less occupational gender segregation. Thus, the benefits of labor market integration for women hold even when the gender composition of their jobs is controlled.

Together, these findings provide strong evidence that the process behind what has been called gender devaluation has a significant local dynamic. Although we agree that women's work roles are devalued by pervasive gender bias, previous research has been unable to account for local variation in devaluation. Our argument is not that devaluation is *either* national or local—it is undoubtedly *both*. However, previous research has focused primarily on either broad, national occupations (e.g., England 1992; Tam 1997) or the organizational context of devaluation divorced from the local labor market (e.g., Baron and Newman 1990; Huffman and Velasco 1997). We argue for the importance of local context, but do not downplay widespread, general processes leading to inequality. Even an apparently universal process must be implemented locally because of the local nature of wage-setting in the United States, and will therefore be prone to local variation. Further research that attends to variation in the wage effect of gender composition, or other processes leading to gender inequality, would add greatly to our understanding how inequality is created and maintained.

How does occupational integration affect the local process of devaluation? There are several ways that integration may influence the devaluation process. Cotter et al. (1997: 715) argue that integration can affect women's wages by changing local norms, increasing pressure on employers and managers to promote women, and boosting the representation of women in decision-making positions (Jacobs 1992). Thus, occupational integration may contribute to the institutional environment (Beggs 1995) in which wages for male- versus female-dominated occupations are determined. Over time, then, "areal customs affecting employer practices" may lead to "legacies of cultural differences" regarding processes that lead to unequal labor market outcomes (Baron 1984:49–50). For example, Cohen and Huffman (2003) analyze establishment-level data and find that gender devaluation is stronger in more segregated labor markets, net of various personnel practices and other demand-side factors (e.g., internal labor markets and establishment size).

On the other hand, it is also possible that occupational integration and gender devaluation are determined by some antecedent metropolitan area-level characteristic, such as the demand for female labor. Cotter et al. (1998) find that labor markets with greater proportions of traditionally female jobs exhibit less gender inequality. Indeed, across metropolitan areas, female labor demand and occupational integration are positively correlated ($r = .47$, $p < .001$). And we do find that there is significantly less within-job gender inequality in markets that tilt toward female-dominated occupational structures (final model, results not shown). However, our findings regarding the effects of integration persist even when the level of demand for female labor is controlled.

Advocates of comparable worth argue that male- and female-dominated jobs should be compensated according to skill and other requirements, countering the tendency for women's work to be broadly devalued in the labor market (England 1992; Nelson and Bridges 1999). Our results confirm that such a policy would do much to address gender inequality in pay. However, comparable worth alone does not address the tendency for men and women to work in segregated occupations. Enforcement of antidiscrimination laws could remedy some aspects of gender segregation; however, to date no significant national policy initiative has directly targeted gender segregation. This is especially important given our finding that spatial variation in gender inequality is tied to the local level of occupational segregation. A policy that reduced occupational segregation—especially in conjunction with a comparable worth approach—could thus reduce the tendency for some labor markets to be more inequitable than others, in addition to reducing the overall level of gender inequality.

Our results suggest several avenues for future research, the first of which concerns these metropolitan area-level mechanisms. Although we have identified the local context as a crucial arena for these processes, additional research is needed to uncover the processes supporting gender devaluation, and how they interact with local labor market dynamics. These processes may be rooted in aspects of the institutional or legal environment, they may be cultural or political, or they may lie in other unmeasured aspects of the workplace or structural context.

Second, although our analysis was able to take advantage of the reach of decennial census data, using these data requires adopting a proxy for jobs. We believe this is the best approximation of local jobs possible, but analysis of workplace-based data would provide a closer test of these hypotheses.

Third, although we control for race/ethnicity at the individual, job, and labor market level, it is possible that there also are racial/ethnic interactions with these gender effects. For example, Black-female devaluation might differ from White-female devaluation (Nelson and Bridges 1999). Introducing these interactions, with the many complications they entail, will require additional theoretical and methodological development, but may prove fruitful for understanding the dynamics of wage inequality for people in diverse jobs and labor market contexts.

All of these suggestions point to the necessity of research on local variation in valuation processes as one important component of the structure of gender inequality.

***Philip N. Cohen*** *is Assistant Professor at the University of California, Irvine. His research involves micro-macro connections in social inequality, including the effects of labor market racial/ethnic composition and inequality within and between jobs across labor markets. His other recent work concerns women's employment, including inequality between Black and White women, occupational segregation and the household division of labor, and the relationship between family structure and inequality within and between families.*

***Matt L. Huffman*** *is Assistant Professor at the University of California, Irvine. His current research examines organizational patterns of race and gender inequality and how they vary across labor markets. Other recent work investigates the role of job seekers' work-related contacts in creating and sustaining the sex segregation of jobs.*

## APPENDIX A

**Variance Components**

Our first model is a fully unconditional, three-level model that includes no independent variables at any level. This model is equivalent to a one-way analysis of variance with random effects (Bryk and Raudenbush 1992) showing how wage variation is distributed across the three levels of analysis. The fully unconditional model can be expressed in three parts, where the subscripts $i$, $j$, and $k$ represent individuals, jobs, and labor markets, respectively:

$$Y_{ijk} = \pi_{0jk} + e_{ijk}, \tag{A-1}$$

$$\pi_{0jk} = \beta_{00k} + r_{0jk}, \tag{A-2}$$

$$\beta_{00k} = \gamma_{000} + u_{00k}. \tag{A-3}$$

In the individual-level model (equation A-1), $Y_{ijk}$ equals the logged earnings of individual $i$ in job $j$ in labor market $k$, and $\pi_{0jk}$ represents the mean earnings in job $j$ in labor market $k$. Last, $e_{ijk}$ is the deviation of person $ijk$'s wage from his or her job-level mean. This term is assumed to be normally distributed with a mean of 0 and constant variance ($\sigma^2$). In the job-level model (equation A-2), each job-level mean can be expressed in terms of variability around the associated labor market mean. Here, $\beta_{00k}$ is the mean wage in labor market $k$, and $r_{0jk}$ represents the deviation of job $jk$ from the labor market mean; these deviations are, by assumption, normally distributed with a mean of 0 and constant variance ($\tau_\pi$). In the labor market model (equation A-3), $\gamma_{000}$ is the grand mean, and $u_{00k}$ is the deviation of labor market $k$'s mean from the grand mean. These deviations are, by assumption, normally distributed with a mean of 0 and constant variance ($\tau_\beta$). It follows that $\sigma^2 / (\sigma^2 + \tau_\pi + \tau_\beta)$ equals the proportion of the variance in wages among individuals within jobs, $\tau_\pi / (\sigma^2 + \tau_\pi + \tau_\beta)$ yields the proportion of the variation in wages among jobs within labor markets, and $\tau_\pi / (\sigma^2 + \tau_\pi + \tau_\beta)$ equals the proportion of wage variation that exists among labor markets.

Table A-1 shows variance components for up to seven terms in each of five models. (All variance components on the table have a two-tailed $p$-value of less than .001.) The fully unconditional model (Model 0) indicates that 65 percent of the total variability in log wages is due to variability in wages among individuals within jobs, 31 percent is due to variation among jobs within labor markets, and the remaining 4 percent is due to differences across labor markets. This decomposition illustrates the

**Table A-1. Variance Components for Hierarchical Linear Models**

| | Model 0 (Unconditional) | | Model 1 (+ Female Dummy) | | Model 2 (Individual Controls[a]) | | Model 3 (Job Controls[b]) | | Model 4 (Full Model[c]) | |
|---|---|---|---|---|---|---|---|---|---|---|
| Variable | Component | Proportion of Total | Component | Proportion of Total | Component | Proportion of Total | Component | Proportion of Total | Component | Proportion of Total |
| *Intercept* | | | | | | | | | | |
| Individual, $\sigma^2$ | .264 | .65 | .252 | .65 | .231 | .72 | .231 | .83 | .231 | .86 |
| Job, $\tau_{\pi 00}$ | .125 | .31 | .119 | .31 | .075 | .23 | .032 | .11 | .032 | .12 |
| Labor market, $\tau_{\beta 00}$ | .017 | .04 | .017 | .04 | .013 | .04 | .014 | .05 | .006 | .02 |
| Total | .405 | 1.00 | .388 | 1.00 | .319 | 1.00 | .277 | 1.00 | .269 | 1.00 |
| *Female ($\pi_1$)* | | | | | | | | | | |
| Job, $\tau_{\pi 1}$ | — | | .018 | .93 | .020 | .92 | .007 | .84 | .008 | .94 |
| Labor market, $\tau_{\beta 10}$ | — | | .001 | .07 | .002 | .08 | .001 | .16 | .001 | .06 |
| Total | | | .019 | 1.00 | .022 | 1.00 | .008 | 1.00 | .008 | 1.00 |
| *Proportion Female, $\beta_{01}$ (Effect on the Intercept)* | | | | | | | | | | |
| Labor market, $\tau_{\beta 01}$ | — | | — | | — | | .006 | 1.00 | .003 | 1.00 |
| *Proportion Female, $\beta_{11}$ (Effect on Individual Female Effect)* | | | | | | | | | | |
| Labor market, $\tau_{\beta 11}$ | — | | — | | — | | .003 | 1.00 | .001 | 1.00 |

[a] Controls include own children, foreign-born, disabled, race/ethnicity, education, potential experience, married, and weekly hours.

[b] Controls include race/ethnic proportions, proportion foreign-born, general educational development, physical demands, standard vocational preparation, and industry.

[c] Controls include population size, unemployment, proportion Black and proportion Hispanic, durable goods manufacturing, female labor demand, region, and net internal migration.

importance of including the job level in the analysis. If we analyze our own data to exclude the job level (not shown), we find 4 percent of the wage variance at the labor market level, and the remaining 96 percent is at the individual level. In the two-level wage models commonly used for testing labor market effects (Cohen 2001; Haberfeld, Semyonov, and Addi 1998; McCall 2001b), variation at this level is not accounted for. Although a small portion of the total variance is observed at the labor market level, it is important to note that the substantive effects of labor market characteristics still may be large. Merlo et al. (2001) provide a discussion of the importance of contextual-level variables even when they account for a small portion of the total variance.

Table A-1 also allows us to decompose the variance in the effects of being female into between-job and between-metropolitan area components (second panel). This shows that between 84 and 94 percent of the variance in the effects of the female dummy variable is observed at the job level, and 6 to 16 percent is observed at the labor market level. Substantively, that means that the female penalty varies more across jobs within labor markets (e.g., between firefighters and lawyers in Detroit) than it does across labor markets (e.g., lawyers in Detroit versus lawyers in Atlanta). However, again, the variance occurring at the labor market level is statistically significant, and as we show above, labor market variables significantly impact the gender gap at the individual level.

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[9] **All Women Benefit: The Macro-Level Effect of Occupational Integration on Gender Earnings Equality**
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