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INDUSTRIAL SHIFTS, FEMALE EMPLOYMENT, AND OCCUPATIONAL DIFFERENTIATION: A DYNAMIC MODEL FOR AMERICAN CITIES, 1960–1970

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Abstract—Sex-linked occupational differentiation has been seen as influenced by both the industrial structure of the economy and the sex composition of the labor force. Here, with a sample of 70 SMSAs, it was found (a) that the odds of men relative to women of joining professional and managerial occupations increased between 1960 and 1970, and (b) that this increase was dependent on the growth of tertiary industries and the greater number of women joining the cash economy. The observed effect of industrial shifts on sex-occupational differentiation, however, is argued to be a spurious consequence of the gender-composition of the work force. Specifically, the development of tertiary industries generates greater demand for female labor. Intensive recruitment of women to the labor force in turn increases occupational differentiation because females, in sex-typed labor markets, are likely to be channeled in disproportionate numbers away from upper-status occupations. The findings demonstrate that traditional modernization theory is unable to account for this. However, the results lend support to expectations derived from a labor market sex-segmentation approach.

One of the most significant aspects of social change since World War II is the dramatic increase in female labor force participation. As a result, demographers and sociologists have devoted greater attention to issues of female employment. This ever-growing literature shows high agreement. Although the proportion of women in the work force has increased considerably, employed women remain concentrated in only a few occupations (e.g., Blau, 1975; Ginzberg, 1968; McLaughlin, 1978; Oppenheimer, 1970; Rosenfeld and Sorenson, 1979). In particular, women are likely to be segregated in low-status, poorly-paid occupations, in jobs of low power and authority, and in marginal industries (e.g., Blau, 1977; Bridges, 1980; Collver and Langlois, 1962; Grimm and Stern, 1974; Oppenheimer, 1970; Semyonov, 1980; Treiman and Terrell, 1975; Wolf and Fligstein, 1979a, 1979b).

While the literature on male-female occupational differentiation has become substantial, only a few studies have examined its trends (Blau, 1977; Gross, 1968; Tsuchigane and Dodge, 1974; Williams, 1975, 1976, 1979). All reported extreme occupational differentiation between the sexes. And most suggest (e.g., England, 1981) that sex-based differentiation is decreasing. Surprisingly, though, no one has explored the mechanisms that generate changes in sex-linked occupational differentiation.

Two theoretically-based explanations exist to account for trends in occupational differentiation by sex. The first centers on long-held ideas regarding modernization's tendency toward universality in the division of labor. Conversely, the second focuses on the growth of particularism which is argued to inhere in the social organization of work. This latter rationale posits that labor markets are sex-segmented and occupations sex-labeled.

The traditional view of the role of modernization in the division of labor traces to Durkheim (1964). This explanation expects the relationship among industrialization, specialization of labor, and ascription to produce decreased differentiation of labor by sex over time. Modern economies are more likely to operate according to universalistic criteria. In particular, with increased industrialization and technologization, the criteria for occupational differentiation shifts from ascriptive to achieved characteristics (e.g., Davis, 1948; Moore, 1951; Parsons, 1951; Treiman, 1970).

Specialization of labor that results from growing industry and higher technology demands more skilled labor. The shift in labor force allocation occurring with industrialization has been from agriculture to services. Singelmann points out that "(m)ost services require little physical strength and this should remove a further barrier to the equal employment of women" (1978a, p. 12). In fact, abilities acquired through individual achievement become paramount for modernization proponents. This approach, then, assumes an increasing need exists to rationally allocate labor to occupational positions based on skill (without regard to gender) as technology becomes more complex and the division of labor more fine. Accordingly, one would expect a reduction in gender-linked occupational differentiation (see Williams, 1975, 1976, 1979).

The opposite view operates on the premise that occupations are sex-typed. The logic expressed by Oppenheimer (1970) and others (e.g., Lloyd and Niemi, 1979; Treiman and Terrell, 1975) suggests that female-typed occupations actually gain more women workers over the years. The rationale embodied in theories of labor market segmentation (e.g., Blau and Jusenius, 1976; Oppenheimer, 1970) leads one to expect that if occupations are sex-typed, as more women enter the labor market they are more likely to be channeled in disproportionate numbers to the female-labeled occupations.

In her classic 1970 study, Oppenheimer examined the patterns of women's occupational concentration, finding that women workers were largely centered in occupations that were predominantly female. From 1900 to 1960, between 60 and 73 percent of working women were in occupations in which over half the workers were female. Further, between 30 to 48 percent worked in occupations comprised of 80 percent or more women. Although she cautions against using her results for trend analysis, her theoretical formulation implies an expectation of increased occupational sex-differentiation. This is due to a series of interrelated changes associated with "post-industrialization" (Bell, 1973; Clark, 1947; Liebenstein, 1957). Greater technologization and industrialization lead to increased specialization of labor and rising per capita income. In response, industrial and occupational shifts take place, expanding job opportunities in tertiary (mainly service) industries. Oppenheimer shows occupations in tertiary industries to be "female-demanding." Hence an increase in the recruitment of women to specific occupations (e.g., female-labeled) is likely to increase differentiation.

Oppenheimer's thesis gains additional support from several studies. For example, Treiman and Terrell (1975) demonstrate that occupations that tend to gain more women workers are those that initially had a high proportion of females and those in which men were poorly paid in previous decades. Their findings suggest that differentiation is not only increasing, but that it also takes a discriminating form. Similar trends are revealed from researchers examining the phenomenon through 1970 and beyond (e.g., Lloyd and Niemi, 1979).

In light of the series of findings and the theoretical point of view expressed above, one can examine Semyonov's (1980) cross-national findings from a longitudinal perspective. He shows that the odds of women joining high-status occupations decrease as the level of economic development grows because such societies recruit more women to their cash economy. This suggests that females, like other subordinate groups, face greater occupational disadvantages as their proportion in the labor force increases. To put this rationale in a dynamic sequence, industrialization results in a greater rate of female work force involvement which, in turn, leads to an increase in occupational differentiation.

Central to either argument is the notion of industrial transformation. Early studies of the subject (e.g., Clark, 1940; Fisher, 1935) pointed to the recession of agriculture and the growth of manufacturing and eventually service industries. Recently in a series of studies (Browning and Singelmann, 1975; Singelmann, 1978a, 1978b; Singelmann and Browning, 1980) it has been shown that sectoral shifts in employment which accompany greater industrialization result primarily in the rapid growth of services (Singelmann, 1978a). This transformation, in turn, has produced a great influx of workers to white collar (service, clerical, professional and managerial) occupations (Singelmann and Browning, 1980).

Thus, while both theses suggest that occupational differentiation between the sexes is affected by changes in the industrial structure, they predict opposing results. Traditional modernization theory proposes that industrialization not only increases opportunities for women to work, but also leads to decreased differentiation. The segmented labor market approach suggests that because work categories are sex-labeled, increased female labor force participation is expected to increase occupational differentiation.

Although neither theory has been tested directly, both can claim indirect empirical support. In the present paper, we seek to discern whether changes in occupational differentiation by sex are related to changes in the sex composition of the labor force as well as to changes in the structure of the economy. First we ask whether differentiation is increasing, decreasing, or stable over time? Second, if change is the case, is it dependent on both the enlargement of tertiary (service) industries and on the number of women who join the economically active labor force? Our analysis centers only on the 1960–1970 decade. Consideration of longer term trends is beyond the scope of this study and is discussed elsewhere (Scott, 1983).

MEASURING OCCUPATIONAL DIFFERENTIATION

Three basic methods have been utilized to arrive at measures of change in occupational differentiation between the sexes: male-female ratios (Tsuchigane and Dodge, 1974), index of dissimilarity (Blau, 1977; Gross, 1968; Williams, 1975, 1976, 1979), and odds ratios extracted from log-linear models (Semyonov, 1980).[1] It is essential to discuss each method's properties in order to delineate their limitations, advantages, and meanings.

To examine these measures, consider a simple cross-classification of occupation by gender. The f_{ij} are the observed frequencies for the ith sex (i = 1, 2; 1 = males, 2 = females) and the jth occupational category (j = 1, 2; 1 = a given occupational category, 2 = all other categories). The proportion of a certain sex in a given occupational category is p_i (specifically, p_1), where $p_1 = f_{i1}/f_{.1}$; and the proportion of a given sex in the labor force is $p_{i.}$, where $p_{i.} = f_{i.}/f_{..}$.

Gross (1968), Williams (1975, 1976, 1979) and Blau (1977) center on the widely used and accepted index of dissimilarity (D). The index of dissimilarity yields a general estimate of the percentage of women (or men) who would have to change occupations in order to achieve equal distributions by sex. If the values of D equal zero, then no difference exists in the occupational distributions of males and females.

The index of dissimilarity can be put in the terms defined above as:

$$D = 100 \cdot \frac{\sum_j \left| \frac{f_{1j}}{f_{1.}} - \frac{f_{2j}}{f_{2.}} \right|}{2}.$$

Here, equality is defined as the situation when

$$\sum_j p_{1j} = \sum_j p_{2j} = \sum_j p_j;$$

in a 2 × 2 table this simplifies to $p_1 = p_2 = p$.

The index of dissimilarity computed in its original form does not control for changes that may be taking place due to shifts in the occupational structure. The greater the size of an occupational category, and the greater the number of categories, the higher the differentiation values (Cortese et al., 1976; Taeuber and Taeuber, 1965). The index, though, can be "size standardized . . . [and] calculated with a fixed number of occupational categories so that both the size and the number of occupations are controlled" (Williams, 1979, p. 78). Size-standardized index of dissimilarity represented in the previously defined terms would be:

Size-Standardized D =

$$100 \cdot \frac{\sum_j \left| \frac{p_{1j}(1000)}{p_{.1}} - \frac{p_{2j}(1000)}{p_{.2}} \right|}{2}$$

Equality for size-standardized D is defined as the condition where

$\sum_j p_{1j}/p_{.1} = \sum_j p_{2j}/p_{.2}$; this reduces in a 2 × 2 table to $p_1/p_{.1} = p_2/p_{.2} = p_{..}$.

Although the size-standardized index of dissimilarity is supposed to resolve this problem, it actually treats each category as if it were the same size. This weighting procedure inflates the impact of small categories as it devalues the influence of larger categories. Unless accurate population parameters are available that suggest sex-segmentation is roughly constant by category size, this adjustment produces a biased estimate.

For some time researchers have known that the index of dissimilarity suffers from severe problems when used to compare populations across time or place (e.g., Cortese et al., 1976; Duncan and Duncan, 1955; Hornseth, 1947; Jahn et al., 1947; Taeuber and Taeuber, 1965). In fact, Gibbs introduced the size standardization adjustment in attempting to render D comparable across populations (1965, pp. 163–164). Cortese et al. point out, however, that parameters such as category size are "*variables* which must be measured and included in any calculation where they vary" (1978, p. 591, emphasis in original). This index accurately captures aggregate differentiation in summary form for a single population at one point in time. But its inability to control for structural changes (i.e., size) in occupations limits its use as a valid longitudinal measure.

Tsuchigane and Dodge (1974) took a different measurement route. They employed sex ratios for each major occupational category. This index is defined as the ratio of the number of males in a given occupational group to the number of females in the same group. The sex ratio expressed in terms of the hypothetical cross-classification becomes:

$$\text{Sex Ratio} = f_{1j}/f_{2j}$$

Equality is defined for the case where $f_{1j}/f_{2j} = 1.00$ or $f_{1j} = f_{2j} = p_1 = p_2$. One should consider that this measure is greatly influenced by the proportion of women in the work force. That is, if more females are joining the labor force, then the ratio between the sexes is expected to decline for this reason alone. Unlike indexes of dissimilarity, this measure controls for the occupational structure. But sex ratios are blind to changes that may occur due to the shifting sex composition of the cash economy.

Both sex ratios and indexes of dissimilarity are problematic. Neither properly controls simultaneously for changes both in the occupational structure and in the sex composition of the labor force. Both are significant determinants of occupational differentiation. Fortunately, a method which enables one to obtain an indicator of occupational differentiation while controlling for both the sex composition and the occupational structure of the labor force is readily available within the log-linear context (Bishop, et al., 1975; Goodman, 1972; Knoke and Burke, 1980).

By using a log-linear model, it becomes possible to extract the sex-occupation interaction for each time point while controlling for the sex and occupational structure of the work force. This interaction represents the relative odds of women (or men) to belong to a given occupational category. Hence the sex-occupation interaction term indicates the rate of sex-occupational differentiation for a given occupational group (Semyonov, 1980). The saturated log-linear model expressed in the previously defined terms is:

$$\ln F_{ij} = \lambda + \lambda_{(i)}{}^{S} + \lambda_{(j)}{}^{O} + \lambda_{(ij)}{}^{SO}$$

where λ is the grand mean, $\lambda_{(i)}{}^{S}$ is the main effect of sex, $\lambda_{(j)}{}^{O}$ is the main effect of occupation, and $\lambda_{(ij)}{}^{SO}$ is the sex-occupation interaction effect. Equality for this formulation in the distribution of occupations across sex becomes represented by the model of independence where the $\lambda_{(ij)}{}^{SO}$ effect is set equal to zero. Setting the log of the cross-products odds ratio equal to 0.00 becomes equivalent, $\ln(f_{11} \cdot f_{22}/f_{21} \cdot f_{12}) = 0.00$.

The log of the cross-products odds ratios, $\ln(f_{11} \cdot f_{22}/f_{21} \cdot f_{12})$, represents a unique estimate of the chances of males ($i = 1$) relative to females ($i = 2$) to belong to a given occupational group ($j = 1$) as opposed to the remaining occupational categories ($j = 2$). Thus positive values indicate odds favoring male employment in a given occupation vis-a-vis the remaining ones, while negative valences show the reverse favoring females. Zero represents equal employment odds for the sexes. Using this scheme, we will examine changes in the odds of the sexes to belong to high-status occupations from 1960 to 1970 across American cities.

ANALYSIS AND FINDINGS

This inquiry is restricted to 70 SMSAs of populations over 250,000 that did not change boundaries during the 1960-1970 period.[2] Data were obtained directly from the 1960 and 1970 publications of the U.S. Census of Population. Three variables were estimated for each city for both 1960 and 1970: (a) an estimate of gender-linked occupational differentiation; (b) size of the tertiary or service industries; and (c) the sex composition of the labor force. A detailed list of the SMSAs and the distribution of the main variables included in the analysis can be found in the Appendix.

Occupational differentiation was estimated using the log-linear technique discussed above. Occupations were dichotomized into professional and managerial employment—the high-status—versus all other work. The dichotomy chosen to illustrate the mechanisms of change in sex-occupational differentiation is an important one. The high-low status distinction represents relative access to desirable rewards including income, respect, and authority. Further, it has been shown that managerial and professional occupations are those growing most rapidly due to advanced industrialization (Singelmann and Browning, 1980, p. 259). This classification also enables one to reduce the data to a series of fourfold tables: high-low-status occupations by sex for each city in each decade. Thus the two separate sets of 2 × 2 × 70 matrices make it possible to derive a unique estimate of the interaction between sex and occupation for each city in 1960 and in 1970.[3] These interaction terms represent a vector of the relative odds (to other cities) of men (or women) belonging to upper-status occupations at each point of time.

In a strict sense, the dependent variable should be referred to as "the relative odds of the sexes to be employed in higher-status occupations compared to all other occupations." But it is too long and cumbersome a label for repeated use. Instead, we prefer to call it sex-linked occupational differentiation while keeping its operational meaning in mind.

The two theoretical determinants of occupational differentiation—industrial structure and sex composition—were also estimated for each city for the two time points. Industrial structure was measured as the proportion of the labor force employed in tertiary industries.[4] The sex composition of the labor force is comprised of the percentage of women in the employed, civilian work force aged 14 or over.

Having estimated these variables for each city in 1960 and 1970, we will first examine changes in their distributions over the decade and the relationship between any pair of variables. Table 1 presents the means, standard deviations and correlations between 1960 and 1970 indicators of the three variables providing a descriptive overview.

Table 1 shows that means of all three variables increased considerably during the decade.[5] While growth in tertiary industries and in the proportion of workers who are women is already well established, the increase in the rate of occupational differentiation is rather a novel finding, though not a novel argument (Oppenheimer, 1970; Rogers and Goudy, 1981; Semyonov, 1980; Treiman and Terrell, 1975). The odds of men joining high-status occupations were higher than for women in 1960, and became even more so by 1970. Even though the percentage of females in high-status occupations has risen, when considering the increasing rate of overall female employment, the actual odds of women working in upper-status occupations have declined. This initially startling finding becomes understandable when cast in the framework of segmented labor market theory.

Looking at the correlation coefficients, one can observe a substantial degree of stability during the decade in the position of cities in the hierarchical scale of the three variables. The strong tendency of cities to maintain their relative position on such hierarchical scales is as expected, especially when social systems are compared over relatively short periods of time. Looked at another way, these coefficients may be regarded as indicators of change rather than stability, since the relative position of many of the cities in 1970 cannot be predicted precisely from their 1960 position. The fact that the correlations are substantially less than $r = 1.00$ can indicate that there is some looseness in the system, and the $1 - r$ can serve as an indicator of change (Collver and Semyonov, 1979). Although the mean of all three variables increased during the decade, apparently some cities grew more in the rate of differentiation, size of tertiary industries, or the proportion of women in the economically active labor force, while others grew less than one would predict from their 1960 values. Later in the anal-

Table 1.—Means, Standard Deviations, and Correlations of the Variables Employed in the Analysis for 70 American Cities: 1960–1970

Variables	Means 1960	Means 1970	Correlations $r_{1960 \cdot 1970}$
Percentage of workers in tertiary industries	52.2 (8.8)	59.5 (8.7)	.927
Percentage of females in the labor force	33.5 (2.9)	38.0 (2.4)	.905
Sex-linked occupational differentiation[a]	.319 (.231)	.368 (.190)	.885

NOTE: Standard deviations in parentheses.

[a]Computed as the relative odds of the sexes working in professional and managerial occupations. For definition of sex-linked occupational differentiation, see text and note 3. Positive values indicate higher odds for men joining high-status occupations.

ysis we will make use of this property of change to explore the underlying dynamics of sex-linked occupational differentiation.

The analysis that follows will examine whether unpredicted changes regarding the three variables are interrelated. Before providing a direct test of the issue, we shall inquire, first, whether the relative odds of the sexes working in upper-status occupations are influenced by the industrial and sex structure of the work force in each of the two decennial census points. Table 2 presents correlation and regression coefficients while predicting occupational differentiation in 1960 and 1970, respectively.

Column 1 of Table 2 displays the cor-

Table 2.—Correlations and Standardized Regression Coefficients Predicting Odds of the Sexes Working in Professional and Managerial Occupations: 70 American Cities, 1960 and 1970

Independent and Control Variables	1960 Correlation (1)	1960 Beta (2)	1960 Beta (3)	1970 Correlation (1)	1970 Beta (2)	1970 Beta (3)
Percentage of workers in tertiary industries	.413	.078 (.115)	-.157 (.162)	.300	.007 (.103)	-.312 (.109)
Percentage of females in the labor force	.630	.586 (.115)	.708 (.119)	.658	.655 (.103)	.771 (.088)
Female median education	.275		.305 (.132)	.400		.505 (.102)
Percentage non-white	.017		-.151 (.105)	-.048		-.196 (.086)
Region (South/non-South)[a]	.133		.079 (.113)	.030		.082 (.090)
R^2		.402	.474		.433	.628

NOTE: Standard errors in parentheses.

[a]South = 1, non-South = 0.

relation coefficients between sex-linked occupational differentiation and the two independent variables as well as three control variables (South/non-South regional dichotomy, percentage non-white, and women's median educational level). In column 2 we let differentiation be a function of the two independent variables while in column 3 we introduce the control variables into the equation.

The results reported in Table 2 hold for both 1960 and 1970. Column 1 of Table 2 indicates that at each time point, both the size of tertiary industries and percentage female in the labor force are significantly related to the relative odds of the sexes to join higher-status work categories. That is, places characterized by larger tertiary economies or larger proportions of employed women are more likely to exclude females from professional and managerial occupations. Tertiary industries and percentage female in the labor force, however, are also related by $r = .570$ in 1960 and $r = .443$ in 1970. Thus in columns 2 and 3 we show the estimated partial regression coefficients while predicting occupational differentiation. The findings suggest that the effect of sex composition intervenes in the effect of industrial structure on differentiation. Apparently, places characterized by larger tertiary industries tend to recruit more women to their cash economies. Increased participation, in turn, results in decreased opportunities for women joining the upper-status occupations.

Although the findings reported so far support the segmented labor market expectation rather forcefully, they bear only indirectly on the dynamics of occupational differentiation. It was suggested, nonetheless, that a longitudinal increase in occupational differentiation is a direct consequence of similar changes in tertiary industries and in the proportion of females in the work force. In order to test such a dynamic model, one has first to overcome the problem of autocorrelation. Since the 1960 and 1970 distributions of each variable are highly correlated (see Table 1), the analysis can center on the residuals (Bohrnstedt, 1969). That is, one has to compute the size of the deviations of the 1970 values from those predicted from their 1960 distributions. Such estimates provide indicators of the extent to which cities grew more (or less) than predicted given their 1960 values regarding: (a) occupational differentiation; (b) tertiary industries; (c) proportion of employed women. These estimates represent the $1 - r$ aggregate measures of change discussed earlier. Table 3 examines the relationships among the residualized variables in order to find whether changes in the relative odds of the sexes to work in professional and managerial occupations across American cities during the 1960–1970 decade were generated by changes in the industrial and sex structure of their labor forces. While column 1 presents correlation coefficients, columns 2 and 3 show standard regression coefficients of the residualized variables.

The findings reported in Table 3 support the segmented labor market thesis. A decrease in women's odds of working in high-status occupations is significantly related to increases in tertiary industries and the proportion of employed women. The residualized differentiation variable is related to the residualized tertiary variable with a correlation $r = .309$, and to the percent women residualized variable with a correlation $r = .565$. The two residualized independent variables are also interrelated by $r = .496$, indicating that unpredicted change in tertiary industries is associated with similar changes in the percentage of females in the labor force.

The regression equation displayed in column 2 of Table 3 provides further support to the argument. The original correlation observed between the tertiary industry residual and the differentiation residual appears to be spurious and thus alternatively interpretable as a primary effect of the increase of the per-

Table 3.—Correlations and Standardized Regression Coefficients for Residualized Variables Predicting Change in Odds of the Sexes Working in Professional and Managerial Occupations: 70 American Cities, 1960 and 1970

	Residual (1960-70) Occupational Differentiation		
	Correlation (1)	Regression (2)	Regression (3)
Residual (1960-70)			
Percentage of workers in tertiary industries	.309	.038 (.116)	-.044 (.127)
Percentage of females in the labor force	.565	.546 (.116)	.576 (.115)
Female median education	.231		.171 (.103)
Percentage non-white	.015		-.149 (.107)
Region (South/non-South)[a]	-.187		-.117 (.111)
R^2		.320	.382

NOTE: Standard errors in parentheses.

[a]South = 1, non-South = 0.

centage of women in the work force. The effect of the percent female residualized variable on the differentiation residualized variable is highly significant (Beta = .546). The control variables included in the regression analysis reported in column 3 did not alter the results observed in column 2. By placing these findings in a dynamic context, one can conclude that during the 1960–1970 decade American cities have experienced considerable growth in tertiary industries. Apparently, these industrial shifts generated intensive recruitment of women to the cash economy, but these women were disproportionately channeled away from higher-status occupations.

CONCLUSIONS

The major objective of this research was to examine changes that occurred in the participation of women in professional and managerial occupations across American cities during the 1960–1970 decade. Although the proportion of women has risen considerably in professional and managerial work, their relative odds to gain high-status occupations have actually decreased. During the 1960–1970 decade, women were shunted away in disproportionate numbers from professional and managerial occupations. The data suggest that growth of tertiary industries is associated with increased participation of women in the cash economy which, in turn, results in increased differentiation.

We have confronted two alternative theoretical models which attempt to explain longitudinal sex-occupational differentiation—the modernization argument versus the labor market sex-segmentation approach. The data indicate that modernization theory is too simplistic to account for the trend. One must understand shifts in occupational differentiation by sex within the context of sex-labeling (see especially the thesis advanced by Oppenheimer, 1970). For women, at least, ascription remains a crucial factor in allocating workers to occupations, especially regarding women's odds to gain employment in professional and managerial positions.

We have focused on the variables central to the theoretical models outlined in the paper. Certainly, the measure of occupational differentiation used here cannot cover all aspects of this phenomenon. In fact, it has been shown elsewhere (Blau, 1975; Kreps, 1976; Lloyd and Niemi, 1979; Patterson, 1973; Quadagno, 1976) that women also tend to be concentrated in lower-status work within the major occupational groupings. Our measure, therefore, is inherently conservative. Although size of tertiary industries and percentage female in the labor force cannot exhaust all possible determinants of differentiation, these two variables provide succinct and meaningful empirical indicators for the theoretical models chosen for this study. Our analysis shows clearly that industrial shifts and the incorporation of women to the cash economy proved to have important consequences for the social organization of labor markets.

NOTES

[1] Autoregression has also been used in this literature (Synder and Hudis, 1976; Treiman and Terrell, 1975). This technique is somewhat complicated by autocorrelation problems (cf., Treiman and Terrell, 1975; p. 183n). More important for this paper, occupational characteristics have been used with this method as the primary explanatory variables. Our focus, instead, is on overriding shifts in the structure of the economy as determinants of differentiation change, rather than on changes in, say, the educational composition of occupations.

[2] The selection of the 70 SMSAs used was not arbitrary. In order to build confidence in the results, only units of analysis could be employed that did not change boundary definitions. Of all the SMSAs of 250,000 or more residents, only 70 did not have their boundaries changed by the Census Bureau between 1960 and 1970. The omission of the remaining SMSAs does not seem to have affected the results. The findings reported in the present paper are consistent with a similar analysis conducted across the 48 contiguous American states. The findings are available from the authors. For a complete listing of the SMSAs utilized in the present paper, see the Appendix.

[3] The saturated log-linear model for the $2 \times 2 \times 70$ occupational contrast for each point of time is:

$$\ln F_{ijk} = \lambda + \lambda_{(i)}^{S} + \lambda_{(j)}^{O} + \lambda_{(k)}^{P} + \lambda_{(ij)}^{SO} + \lambda_{(ik)}^{SP} + \lambda_{(jk)}^{OP} + \lambda_{(ijk)}^{SOP}$$

where λ is the grand mean, $\lambda_{(i)}^{S}$ is the main effect of sex, $\lambda_{(j)}^{O}$ is the main effect of occupation, $\lambda_{(k)}^{P}$ is the main effect of SMSA, $\lambda_{(ij)}^{SO}$ is the sex-occupation interaction effect, $\lambda_{(ik)}^{SP}$ is the sex-SMSA interaction effect, $\lambda_{(jk)}^{OP}$ is the occupation-SMSA interaction effect, and the three-way interaction effect of sex-occupation-SMSA is represented by $\lambda_{(ijk)}^{SOP}$.

[4] The percentage of the employed civilian labor force 14 and over working in tertiary industries comprises workers with jobs in the following: wholesale and retail trade, finance, insurance and real estate, business and repair service, personal services, entertainment and recreation services, professional and related services, and public administration. This classification is provided by the Bureau of the Census in Report by States, Table 186 for 1970 and Table 127 for 1960. From an empirical point of view, tertiary industries could easily be replaced by the percentage of workers employed in manufacturing industries, since these two variables are inversely associated by a correlation exceeding $r \simeq -.9$ in both 1960 and 1970.

[5] It is also interesting to note that standard deviations decreased over the decade. On the average, therefore, cities not only increased in their mean values, but they also became more similar regarding the three variables.

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Appendix Table A.—Percentage Female in the Labor Force, Percentage of the Labor Force Employed in Tertiary Industries, and Sex-Linked Occupational Differentiation for 70 SMSAs, United States, 1960 and 1970

	1970			1960		
SMSA[a]	% Female in the Labor Force	% Employed in Tertiary Industries[b]	Sex-Linked Occupational Differentiation[c]	% Female in the Labor Force	% Employed in Tertiary Industries[b]	Sex-Linked Occupational Differentiation[c]
Albany-Schenectady-Troy, NY	38.8	63.4	.460	34.5	51.3	.411
Albuquerque, NM	38.1	77.2	.585	33.0	68.2	.570
Allentown-Bethlehem-Easton, PA-NJ	38.0	42.4	.542	33.6	35.5	.444
Atlanta, GA	40.2	63.1	.563	36.7	56.0	.623
Augusta, GA-SC	40.7	55.9	.256	37.0	51.5	.202
Austin, TX	40.9	77.1	.559	37.9	70.3	.500
Bakersfield, CA	34.3	60.2	.049	29.5	53.0	.088
Baton Rouge, LA	37.4	65.5	.275	34.9	60.4	.468
Beaumont-Port Arthur, TX	32.5	52.1	-.035	29.4	47.1	.001
Buffalo, NY	36.9	54.4	.234	31.0	43.6	.130
Canton, OH	34.8	46.3	.172	29.3	40.3	.095
Chattanooga, TN-GA	39.4	50.3	.342	34.6	46.9	.373
Chicago, IL	38.5	54.8	.501	33.9	45.3	.521
Columbia, SC	41.3	65.2	.383	39.3	61.2	.259
Denver, CO	39.4	66.4	.557	34.5	58.6	.539
Des Moines, IA	41.1	67.2	.521	36.6	59.6	.470
Detroit, MI	35.4	52.1	.391	30.2	44.5	.361
Duluth-Superior, MN-WI	35.7	59.0	.115	30.6	50.7	.014
El Paso, TX	39.2	65.7	.307	36.1	60.2	.089
Erie, PA	35.3	46.4	.283	31.3	40.9	.082
Fort Lauderdale-Hollywood, FL	38.9	66.8	.509	34.8	61.4	.554
Fort Wayne, IN	38.8	53.4	.611	32.8	47.6	.504
Fort Worth, TX	36.8	54.6	.541	32.5	51.0	.518
Fresno, CA	35.3	63.7	.196	29.4	51.7	.134
Gary-Hammond-East Chicago, IN	32.4	43.3	-.119	26.0	34.2	-.198
Honolulu, HI	41.9	71.4	.319	36.5	60.8	.125
Huntington-Ashland, WV-KY-OH	33.6	50.7	-.026	30.1	46.0	-.039
Jacksonville, FL	40.7	68.3	.392	36.2	61.3	.488
Jersey City, NJ	39.7	50.0	.299	34.9	41.3	.257
Johnstown, PA	34.6	47.9	-.078	29.0	41.9	-.221
Knoxville, TN	37.4	56.9	.364	33.1	49.0	.254
Lancaster, PA	37.9	43.2	.362	34.1	37.7	.284
Lansing, MI	38.6	63.0	.376	32.8	54.2	.265
Las Vegas, NV	35.8	78.1	.260	33.2	68.6	.322
Lorain-Elyria, OH	33.0	50.2	.068	26.9	36.3	-.141

Appendix Table A.—Continued

Louisville, KY-IN	37.8	52.7	.209	32.6	47.7	.301
Madison, WI	41.9	73.2	.540	36.5	61.9	.408
Miami, FL	41.3	65.5	.581	36.1	61.1	.524
Minneapolis-St. Paul, MN	40.4	61.6	.673	35.9	54.2	.560
New York, NY	39.0	65.2	.468	35.0	54.7	.555
Newark, NJ	38.7	55.0	.633	33.5	46.3	.657
Newport News-Hampton, VA	38.5	52.7	.358	33.1	54.3	.180
Norfolk-Portsmouth, VA	39.5	68.2	.165	35.5	62.6	.141
Oklahoma City, OK	39.8	69.8	.603	34.5	63.9	.498
Omaha, NE-IA	39.9	63.3	.361	34.7	52.9	.255
Orlando, FL	38.7	64.5	.499	34.2	55.5	.532
Paterson-Clifton-Passaic, NJ	37.4	54.0	.655	31.4	44.4	.671
Philadelphia, PA-NJ	38.0	56.5	.491	33.6	46.7	.512
Phoenix, AZ	38.5	62.7	.479	40.8	71.3	.886
Pittsburgh, PA	34.0	54.4	.274	28.8	44.6	.098
Portland, OR-WA	38.8	61.7	.449	34.5	55.0	.343
Reading, PA	39.5	42.7	.431	35.1	36.9	.387
San Bernardino-Riverside-Ontario, CA	36.5	62.8	.220	31.8	56.3	.100
San Diego, CA	38.4	67.9	.409	33.0	56.5	.407
San Jose, CA	36.3	56.3	.654	31.0	49.1	.497
Santa Barbara, CA	39.1	70.5	.415	33.0	56.7	.370
Seattle, WA	37.3	60.5	.486	32.9	52.0	.543
Shreveport, LA	40.5	62.1	.300	37.1	59.2	.457
Spokane, WA	38.5	70.8	.323	34.8	62.2	.328
Stockton, CA	34.1	61.0	.068	28.4	50.5	-.170
Syracuse, NY	38.4	58.3	.396	33.5	46.5	.323
Tacoma, WA	37.5	64.2	.212	33.4	57.2	.194
Tampa-St. Petersburg, FL	40.2	65.7	.378	35.3	57.2	.387
Trenton, NJ	40.1	60.4	.624	35.6	51.0	.413
Tucson, AZ	37.0	69.3	.271	33.2	61.2	.235
Tulsa, OK	37.1	57.8	.611	32.7	50.6	.568
Utica-Rome, NY	38.1	54.7	.371	34.3	47.9	.265
West Palm Beach, FL	40.1	62.5	.509	35.4	59.6	.467
Wilkes-Barre—Hazelton, PA	41.3	45.7	.406	38.9	40.5	.282
Youngstown-Warren, OH	33.8	45.8	.094	27.8	39.1	-.050

[a]The Standard Metropolitan Statistical Areas comprise all those with populations of 250,000 or more residents whose boundary definitions were not changed by the Census Bureau between 1960 and 1970.

[b]See Note 4 for definition of tertiary industries.

[c]Computed as the natural logarithm of the odds ratio of the sexes belonging to high- or low-status occupations. Positive values indicate higher odds for men joining high-status occupations. (See text for further details.)