

Models for Describing the Underlying Structure of Sex Segregation¹

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This article introduces a structural approach to analyzing sex segregation data that rests on margin-free measures of the underlying association in sex-by-occupation arrays. The starting point for the analyses is a log-multiplicative model that is formally consistent with the conventional practice of summarizing cross-national variability in a single parameter pertaining to the overall strength of sex segregation. Under this baseline specification, the segregation regime is forced to take on the same basic shape in each country, with the only form of permissible variability being a uniform compression or expansion of the peaks and valleys characterizing the shared segregation profile. Although the latter model does not account for the cross-national variability in our illustrative data, it can be readily generalized in ways that both improve the fit and yield new insights into the structure and sources of sex segregation. These elaborated models can be used to examine the hierarchical structure of segregation, to identify the dominant “segregation profiles” in industrial countries, and to parse out the net residue of segregation at multiple levels of analysis.

The study of occupational sex segregation appears to be entering its take-off period. This can be seen, for example, in the recent resurgence of interest in describing how the structure of sex segregation varies across

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nations, over time, and among industries, organizations, or economic sectors (e.g., Reskin 1993; Brinton and Ngo 1991, 1993; Hakim 1992; Presser and Kishor 1991; Baron, Mittman, and Newman 1991; Jacobs 1989*a*, 1989*b*; Tienda and Ortiz 1987; Tienda, Smith, and Ortiz 1987; Bianchi and Rytina 1986; Bielby and Baron 1984). What makes this takeoff period so distinctive is that it is unfolding *without* any major methodological innovations of the kind that have historically played transformative roles in other subfields of stratification research. The methods deployed by sex segregation researchers have, in fact, remained largely unchanged over the last 30 years, with the index of dissimilarity (D) and its not-so-distant cousins still playing a featured role. At regular intervals, the relative merits of competing indices are ritually debated (see, e.g., James and Taeuber 1985; Massey and Denton 1988; Coulter 1989; Watts 1992), yet such commentary appears to have little influence on the subsequent conduct of sex segregation research. We see no evidence, moreover, of an emerging interest in modeling segregation data; indeed, whereas stratification researchers in other subfields have long since abandoned D (and other indices) in favor of model-based “structural parameters” (see Duncan 1984, p. ix; Sobel, Hout, and Duncan 1986), this standard methodological transition has not yet occurred among sex segregation researchers.

It is not unusual for research subfields to bear the imprint of the time period in which they initially became popular. In the case of segregation research, the first burst of activity occurred precisely when D was emerging victorious from a “ten-year index war” (see Peach 1975, p. 3), and we might therefore expect this index to have shaped the development of segregation theorizing and research in decisive ways. As it turns out, the index of dissimilarity and its close analogues quickly became entrenched in the field, so much so that the rise of log-linear modeling went largely unnoticed despite its (seemingly obvious) relevance to tabular analyses of sex segregation. The purpose of the present article, then, is to document some of the advantages of log-linear and log-multiplicative models in describing, comparing, and explaining patterns of occupational sex segregation. After introducing a general multiplicative framework for modeling sex-by-occupation tables, we will derive a new scalar index of sex segregation and specify the conditions under which such an index can satisfactorily represent cross-national variability. We will also introduce simple multiplicative models that can account for the sex-by-occupation

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association in terms of exogenous variables defined at the level of occupations.

MARGINAL DEPENDENCIES IN SEGREGATION INDICES

It may be instructive to briefly review the methodological state of affairs among sex segregation researchers. As we have already noted, the starting point for most analyses is the index of dissimilarity, where this is defined as

$$D = \sum_{j=1}^J |(F_j/F) - (M_j/M)| \times 100 \times 1/2. \quad (1)$$

We have followed convention (see Duncan and Duncan 1955) in using J to refer to the total number of occupations, M_j and F_j to refer to the number of men and women in the j th occupation, and M and F to refer to the number of men and women in the labor force as a whole. The value of D can be interpreted, therefore, as the percentage of the labor force that must change occupations to bring about a perfect correspondence between the sex ratio within each occupation and the overall rate of female labor force participation (see Winship [1977, p. 1061] for further details and qualifications).

The flaws of D are certainly well known (e.g., James and Taeuber 1985), yet it seems that the full implications of the known have not been sufficiently appreciated. What must of course be stressed is that D is not invariant under multiplicative transformations of the occupational margins; as a result, it becomes difficult to interpret cross-national or temporal variability in D , since the driving force behind such variability can be simple distributional differences in the occupational structure as well as real heterogeneity in the sex composition of occupations (cf. Butler 1987; Silber 1989). It is here, then, that the heritage of D as an index of *residential* segregation reveals itself in an unfortunate way. Although most scholars treat occupational and residential segregation as direct analogues, it is important to recognize that the former is measured with categories that are standardized across the units of comparison (i.e., occupations), whereas the latter is typically calculated in terms of "tracts" or "districts" that are defined in highly idiosyncratic ways.² It is clearly

² In carrying out temporal comparisons, the logic underlying the analysis of occupational and residential segregation is much the same, since in both cases the categories of the classification system (i.e., occupations or census tracts) are largely unchanged over time, thereby making it possible to map each of the categories within the benchmark period into one, and only one, of the categories appearing subsequently. The same type of one-to-one mapping is also feasible when one compares occupations across spatial units (e.g., cities, nation-states), but not when one compares census tracts or districts across such units.

OCCUPATIONAL COMPOSITION

		Margin Dependent	Margin Free
L A B O R	Margin Dependent	Isolation Index; Interaction Index; Variance Ratio Index	Size-Standardized Dissimilarity Index
	Margin Free	Atkinson Index; Gini Index; Index of Dissimilarity; Lieberson's Diversity Index	Odds Ratios (or functions of them)

FIG. 1.—Assorted segregation measures cross-classified by two forms of marginal dependence. The interaction and isolation indices are often denoted xP^*y and xP^*x (see Lieberson 1981). The variance ratio index has been labeled S (Zoloth 1976), η^2 (Duncan and Duncan 1955), and the “revised index of isolation” (Bell 1954). The diversity index cited here was introduced by Lieberson (1969), while the Atkinson index was recently reviewed by James and Taeuber (1985). See Coulter (1989) for a comprehensive review of related indices.

impossible to specify a one-to-one correspondence between the census tracts of different cities, and consequently the incentive for devising margin-free measures is somewhat reduced. In this regard, it would be unreasonable to expect the conventional measures of intergroup inequality to be margin free, since most of them were devised with the empirical case of residential segregation in mind (see fig. 1).

We do not mean to imply that contemporary segregation research is invariably carried out with D alone. In fact, there is a small industry of research based on the premise that marginal effects *should* be purged from the data, with the point of departure typically being some type of modified or corrected version of D . We are referring, for example, to the well-known proposal of Blau and Hendricks (1979) to decompose changes in D into components attributable to occupational restructuring and residual “shifts in sex composition” (p. 199; see also Fuchs 1975; England 1981; Handl 1984; Tienda and Ortiz 1987; Beller 1984; Bianchi and Rytina 1986). In more recent work, Abrahamson and Sigelman (1987) sought to purge D of marginal dependence by regressing it on the “structural propensity toward occupational segregation” (p. 591), while Bridges (1982) proposed to adjust D “based on a comparison of the observed

level of occupational segregation with that expected given the occupational mix" (p. 278). We would interpret work of this kind as expressing an incipient interest in margin-free measures of association; however, given that D has become so entrenched in the field, this interest was inevitably channeled into purely reformist efforts (i.e., modifying D).

The latter approaches have attracted some attention (see, e.g., Stafford and Fossett 1989), but it would appear that the size-standardized index of dissimilarity (D_s) is gradually becoming the de facto standard for comparative analyses of sex segregation (see Gibbs 1965; Gross 1968). We can calculate D_s in the following way:

$$D_s = \sum_{j=1}^J \left| \left[\frac{F_j/T_j}{\sum_{j=1}^J (F_j/T_j)} \right] - \left[\frac{M_j/T_j}{\sum_{j=1}^J (M_j/T_j)} \right] \right| \times 100 \times 1/2, \quad (2)$$

where T_j refers to the total number of males and females in the j th occupation (i.e., $T_j = M_j + F_j$). As indicated in equation (2), D_s will be unaffected by the shape of the occupational distribution, since it standardizes each of the J occupations to the same size. The two numerators in this equation (i.e., $[F_j/T_j]$ and $[M_j/T_j]$) index the female and male proportions in the j th occupation, while the corresponding denominators calibrate these values against the proportions prevailing in other occupations. As the amount of sex segregation increases, the difference between the "calibrated proportions" grows large, and the value of D_s increases in tandem. This type of standardization has been recently applied by Brinton and Ngo (1991, 1993), Presser and Kishor (1991), Williams (1979), Jacobs (1989a, 1989b), and Jacobs and Lim (1992).

It turns out that standardizing D is far from cost free. While the usual standardization does eliminate one form of marginal dependence, it has the perverse effect of introducing a new dependence on the rate of female labor force participation.³ The appeal of D has long been its "scale invariance" (James and Taeuber 1985, pp. 15–17); that is, the value of D is unaffected by simple multiplicative transformations of the sex ratio, and consequently it can safely be used to compare countries, cities, or time periods with differing rates of female labor force participation. However, the same property does not hold for D_s , with the implication being that this type of standardization is seriously flawed for purposes of comparative research. We might well regard D_s as the analogue to the "Rogoff index" of mobility research (Rogoff 1953; see also Hauser 1978), since it suffers from the same types of marginal dependence that doomed

³ This can be readily demonstrated by example. If the number of females in each occupation is multiplied by an arbitrary constant, the value of D_s will typically change.

this index and ultimately ushered in a new era of log-linear modeling. The flaws of D_s have *not* been sufficiently appreciated in the literature to date; if anything, it appears that D_s is becoming increasingly popular, with several recent studies featuring it as one of the key indices of sex segregation (e.g., Jacobs 1989a).

The foregoing comments suggest that the current state of affairs is less than ideal. As indicated in figure 1, the index of dissimilarity is invariant under multiplicative transformations of the sex ratio but not under multiplicative transformations of the occupational margins. At the same time, the size-standardized index successfully eliminates the latter dependence, but only at the cost of losing the scale invariance that characterized the original index. If we wish to eliminate both forms of marginal dependence simultaneously, we have no alternative but to use measures that are functions of cross-product ratios (see Goodman 1991; Becker and Clogg 1989; Altham 1970). Although a few enterprising scholars have already applied log-linear models to sex segregation tables, none of the efforts to date has exploited the full potential of this approach (see Handl 1984; Willms 1982; Semyonov 1980; Semyonov and Scott 1983; Stolzenberg and D'Amico 1977). We think that further analyses based on margin-free measures of sex segregation may lead to nontrivial revisions of our understanding of cross-national and over-time variability in gender stratification.⁴

SCALAR MEASURES OF SEX SEGREGATION

We ought not forget that segregation indices are merely scalar summaries of complex "segregation curves" representing the sex composition of all occupations. To be sure, most segregation scholars are quick to concede that "no single measure is correct for all purposes" (e.g., Lieberson 1980, p. 253; Abrahamson and Sigelman 1987), but this ever-popular disclaimer is a poor substitute for satisfactory descriptive work. It should be recalled that Duncan and Duncan (1955) found D to be an acceptable index of racial segregation only because there appeared to be a "characteristic form for the segregation curves of most large American cities" (p. 214). While it is conventional to assume that the corresponding curves for sex segregation data are likewise invariant, such an assumption has not yet been verified in any rigorous way (but see Roos 1985, pp. 38–66; Gaskin 1979). At this relatively early point in the development of sex

⁴ The comparisons based on margin-dependent measures will be especially misleading when carried out across countries or time periods that differ substantially in occupational composition or in rates of female labor force participation.

segregation research, we would thus agree with James and Taeuber (1985) that a “prudent analyst would construct and visually compare segregation curves” (p. 26; see also Brinton and Ngo [1991] for a related argument). We will be presenting such curves throughout the following analyses.

There is good reason to believe that the distinctive cultures and political histories of advanced industrial countries can live on in ways that affect the contours of their segregation profiles. Among scholars who study detailed patterns of sex segregation, it is now commonplace to cite the integration of Soviet women into medical and engineering positions (e.g., Roos 1985; Blekher 1979; Dodge 1971) or to note that pharmacy and dentistry are female-dominated occupations in Finland, Poland, and Hungary (e.g., Safilios-Rothschild 1976; see also Szelényi and Poster 1991). We will be carrying out analyses designed to reveal whether institutional differences of a more fundamental kind can also generate cross-national variation at the level of major occupational groupings. It has long been argued, for example, that Swedish women are disproportionately channeled into a state-supported service sector that is dominated by traditionally sex-typed jobs (see Ruggie 1984; Scriven 1984). By contrast, the well-known internal labor markets of Japan have the apparent effect of pushing women into blue-collar production work (see Brinton 1988; Kalleberg and Lincoln 1988; Clark 1979), while the traditional cultural and institutional environment of Switzerland encourages women to choose occupations that are compatible with their marginal forms of labor market attachment (see Charles and Buchmann 1994; Buchmann and Charles, in press). The structure of these nation-specific sex segregation systems has been reviewed in more detail elsewhere (see, e.g., Charles 1990); at this point, we merely wish to raise the possibility that some forms of cross-national variability may be revealed in *how* male and female work is separated, and not merely in the degree of such separation. If this is indeed the case, then the detailed contours of sex segregation cannot be adequately described with scalar indices, nor can satisfactory explanatory models be devised when these indices are applied.⁵

A GENERAL LOG-MULTIPLICATIVE APPROACH

We will proceed by fitting a series of association models that are consistent with the conventional practice of summarizing cross-national vari-

⁵ The latter point may be obvious to those familiar with cross-national variability in gender stratification systems. We are merely commenting on the apparent disjuncture between our substantive understanding of such systems and conventional methods for describing them (see Brinton and Ngo 1993).

ability in a single parameter. The following model will serve as our starting point:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{(\phi_k Z_i v_j)}, \quad (3)$$

where i indexes sex, j indexes occupation, k indexes country, m_{ijk} is the expected frequency in cell (i, j, k) , α_k is the grand mean in the k th country, β_{ik} is the country-specific marginal effect for the i th gender, γ_{jk} is the country-specific marginal effect for the j th occupation, ϕ_k is the multiplicative shift effect for the k th country, Z_i is an indicator variable for gender (i.e., $Z_1 = 0$ and $Z_2 = 1$), and v_j is the scale value for the j th occupation.⁶ The distinctive feature of this model is that it scales both occupations and countries *without* assuming any prior ranking (see Goodman 1979a, 1979b, 1981a; Clogg 1982; Xie 1992).⁷ As indicated above, the sex-by-occupation association is expressed in a set of J column effects (v_j), while any cross-national variability in sex segregation has to be absorbed by a set of K multiplicative shift effects (ϕ_k).⁸ If this specification fits the data, it follows that the segregation profile is invariant and that ϕ_k can be used to represent cross-national differences in the underlying “strength” of sex segregation. We will not be following the common practice of simply assuming that a scalar index suffices; instead, we have embedded this assumption in a testable model, with the viability of ϕ_k as a segregation index thus resting on the usual criteria of model fit. If our multiplicative shift model fails to provide a satisfactory fit, we will have to modify it by permitting the shape of the sex segregation profile to vary across countries.

The full set of parameters in equation (3) cannot be uniquely estimated. In the following analyses, we will identify the column effects by con-

⁶ We will be using the same indicator variable (Z_i) for most of the models introduced in this article. The metric that we have used to scale gender is of course arbitrary and inconsequential; that is, not only are the column effects estimated here reproduced perfectly by all scalings in which $Z_2 - Z_1 = 1$, but they can be easily recovered (by multiplying through by c) for all scalings in which $Z_2 - Z_1 = c$.

⁷ We have found it convenient to parameterize the sex segregation profile in terms of column effects. However, given that our classification includes only two rows of data, this specification places no within-country constraints on the sex-by-occupation association. Although there are any number of parameterizations that might be adopted in this context, we think it is elegant and analytically revealing to choose one that represents both countries and occupations as scalable quantities. It should nonetheless be kept in mind that many of the log-multiplicative models introduced below could be rewritten as seemingly simpler log-linear models.

⁸ Whereas prior analysts (e.g., Yamaguchi 1987) have parameterized shift effects in additive form, we have followed Xie (1992) in adopting a simple multiplicative specification (see also Clogg 1982; Becker and Clogg 1989; Fukumoto and Grusky 1992). The advantage in doing so is that the resulting model is invariant under all possible reorderings of the column and level categories (see Yamaguchi 1987, p. 484).

straining them to sum to zero and the marginal and multiplicative shift effects by constraining the parameters for the first row, column, or level to equal one:

$$\sum_{j=1}^J v_j = 0,$$

and (4)

$$\beta_{1k} = \gamma_{1k} = \phi_1 = 1.$$

The following closed-form results hold when the model represented by equation (3) fits perfectly:

$$\ln(\beta_{2k}) = 1/J \times \left[\sum_{j=1}^J \ln(F_{jk}/M_{jk}) \right],$$

and

$$\begin{aligned} \phi_k v_j &= \ln(F_{jk}/M_{jk}) - \left[1/J \times \sum_{j=1}^J \ln(F_{jk}/M_{jk}) \right] \\ &= \ln(F_{jk}/M_{jk}) - \ln(\beta_{2k}). \end{aligned} \tag{5}$$

These results indicate that the main effect of gender for the k th country is merely the mean of the logged sex ratios, while the adjusted column effects ($\phi_k v_j$) are simply occupation-specific departures from that mean.⁹ If the model fails to fit perfectly, the estimates of $\ln(\beta_{2k})$ and $\phi_k v_j$ can of course be recovered by replacing F_{jk} and M_{jk} with their expected values.

THE SEX SEGREGATION DATA

We will apply this general model to an eight-nation data set collected by the International Labour Office (ILO). The ILO classifies labor force members into the following major occupations: (1) professional, technical, and related workers, (2) administrative and managerial workers, (3) clerical and related workers, (4) sales workers, (5) service workers, (6) production and related workers, and transport equipment operators and laborers, and (7) agricultural, animal husbandry, and forestry workers,

⁹ The reader should be reminded that $M_{jk} = m_{1jk}$ and that $F_{jk} = m_{2jk}$. We have maintained our earlier notation (see eqq. [1] and [2]) because it emphasizes the connection between conventional segregation indices and the log-multiplicative measures introduced here.

and fishermen and hunters.¹⁰ We will eliminate the seventh category from the following analyses because of substantial cross-national discrepancies in the procedures for assigning women to agricultural labor (see ILO 1986, p. 3). The final sample counts for each country are presented in appendix A (for more details, see ILO [1985, 1986, 1987]).

The available evidence suggests that the rank ordering of countries on standard segregation indices remains roughly the same under both aggregated and disaggregated occupational classifications (see Jonung 1984; Charles 1990; Jacobs and Lim 1992). However, the standard indices may not fully reveal the potentially distorting effects of aggregation, and we will therefore be carrying out a series of supplementary analyses with disaggregated census data from the United States and Japan (see U.S. Bureau of the Census 1984, table 276, pp. 166–75; Statistics Bureau of Japan 1984, pp. 586–620). We have recoded the latter data into an aggregated version of the detailed 1968 International Standard Classification of Occupations (ISCO), with the aggregations being introduced whenever the original occupational codes could not sustain the more detailed classification.¹¹ The end result is the 45-category classification presented in appendix B.¹²

CONSTRUCTING A MARGIN-FREE INDEX OF SEX SEGREGATION

We will begin our analyses with a global test of cross-national variability in sex segregation. It could well be argued that a basic family resemblance in segregation regimes is generated by cross-nationally shared cultural, economic, and institutional forces. We are referring, in particular, to the worldwide diffusion of a family structure in which women have primary responsibility for childrearing, cooking, and maintaining the home. According to some theorists, women in such families should not only have a diminished incentive to invest in work-related human capital (given that investments must often be repaid over a shorter work life) but should also prefer occupations in which wage depreciation is minimized during

¹⁰ These major categories are based on the 1968 International Standard Classification of Occupations (ISCO), but for convenience we shall refer to them as “ILO categories.”

¹¹ We are indebted to Harry Ganzeboom for his recoding of the Japanese data into the ISCO classification.

¹² Although the following analyses will be based exclusively on sex segregation tables, it would be equally instructive to analyze other forms of segregation data with our general multiplicative approach (e.g., racial occupational segregation). Indeed, whenever there is a one-to-one correspondence between the categories being compared (e.g., occupations), the analyst will usually wish to apply measures that are margin free.

TABLE 1
 LOG-LINEAR AND LOG-MULTIPLICATIVE ASSOCIATION MODELS
 APPLIED TO EIGHT-NATION DATA SET

Model or Contrast	L^2	df	L_h^2/L_i^2	Δ
Model:				
1. Conditional independence ($O \times N + S \times N$)	7,121,442	40	100.0	18.1
2. Constant sex segregation ($O \times N + S \times N + C$)	789,207	35	11.1	4.1
3. Multiplicative shift effect ($O \times N + S \times N + C + A \times N$)...	527,952	28	7.4	2.4
4. Segregation profiles ($O \times N + S \times N + C \times P$)	126,389	25	1.8	1.3
Contrast:				
1. Total variability (model 2)	789,207	35	100.0	
2. Explained variability under model 3 (model 3 vs. model 2)	261,255	7	33.1	
3. Explained variability under model 4 (model 4 vs. model 2)	662,818	10	84.0	

NOTE.— O = occupation; N = country; S = sex; C = column effects; A = global row-by-column association parameter; P = tripartite partition of countries. The index of dissimilarity (Δ) can be interpreted as the percentage of cases that would have to be reallocated to bring the observed and expected values into perfect correspondence.

the many employment interruptions that their family obligations require (see Polachek and Siebert 1993; Polachek 1979, 1981; cf. England et al. 1988). The foregoing arguments suggest that men in all countries should dominate jobs that entail a substantial commitment to the labor market, whereas women should be disproportionately concentrated in jobs that neither require large amounts of human capital nor strongly penalize employment interruptions. Although the traditional family structure may therefore have a powerful standardizing effect on gender stratification systems, we would also cite the cultural diffusion of occupational gender labels and gender-based stereotypes as potential forces for convergence.¹³

The overall effect of such forces can be assessed by fixing ϕ_k at one (for all k) and thereby forcing all countries to share the same segregation profile. As shown in table 1 (see model 2), we find a substantial amount of cross-national similarity in our eight-nation data set, with the model of constant segregation misallocating only 4.1% of the cases and account-

¹³ The occupational categories deployed here are so highly aggregated that one would expect substantial within-category variability on such dimensions as human capital requirements, wage depreciation, and the like. This heterogeneity greatly reduces our leverage in adjudicating between competing arguments about the sources or causes of occupational segregation.

ing for 88.9% of the association under the model of conditional independence. The column effects under model 2 take on a largely expected form; namely, women are underrepresented in managerial and production occupations, and they are overrepresented in clerical, sales, and service occupations (see fig. 2)¹⁴ This evidence of cross-national convergence might be seen as the analogue within the field of sex segregation to the Featherman-Jones-Hauser finding that “the genotypical pattern of mobility . . . in industrial societies with a market economy and a nuclear family system is basically the same” (Hauser and Featherman 1977, p. 16; Featherman, Jones, and Hauser 1975).

At the same time, it is important to explore the cross-national variability in the data, since the model of constant sex segregation clearly fails to fit. The first step in doing so is to ask whether a more limited form of convergence might hold; that is, we can entertain the provisional hypothesis that the underlying segregation profile takes on the same basic shape in each country, while the overall *degree* of segregation is cross-nationally variable. This hypothesis leads us to our full log-multiplicative model (see eq. [3]) in which the peaks and valleys of a generic segregation profile are compressed or expanded in accord with a simple shift effect.¹⁵ The latter model does indeed reveal considerable cross-national variation; as indicated in figure 3, the peaks and valleys for the Swedish curve are 2.7 times more extreme than those for the relatively flat Japanese curve, whereas the results for the six remaining countries fall somewhere between these two poles. However, we would not want to take these estimates too seriously, since our baseline model does not adequately account for the cross-national variability in the data. Under a standard decomposition of the test statistic (for model 2 of table 1), only 33.1% of the total variability can be explained with multiplicative shift effects, while the remaining variability arises from cross-national differences in the structure of the segregation profile itself (see table 1, contrast 2).

The latter result indicates that ϕ_k cannot satisfactorily represent the structure of cross-national variability. Moreover, ϕ_k is not necessarily the best of all possible scalar indices, since it improperly assumes a common segregation profile. If we insist on defining a scalar index in such circumstances, we would be well advised to base it on a model that fits the

¹⁴ The pattern of sex segregation revealed here may not be entirely surprising, but it is nonetheless difficult to reconcile it with the simple hypothesis that women should be disproportionately represented in occupations that require only limited investments in human capital. The latter hypothesis is neither consistent with the negative scale value for production workers nor with the gender neutral scale value for professionals.

¹⁵ The log-multiplicative models presented here were estimated with modified GLIM programs that were based on and inspired by similar programs prepared by Mark P. Becker and Yu Xie.

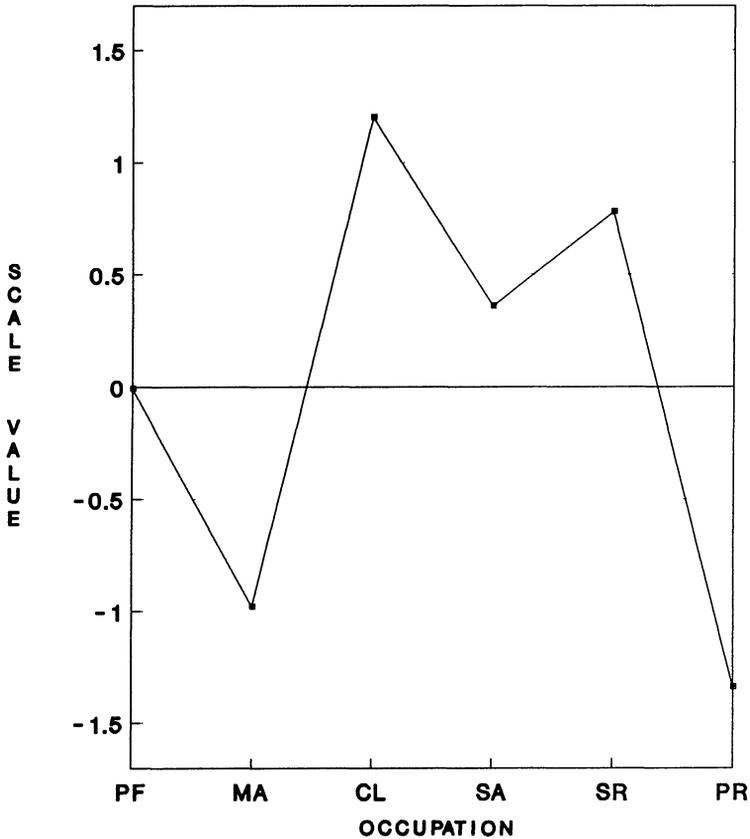


FIG. 2.—A pooled profile of sex segregation. The parameter estimates are taken from model 2 (table 1). In the present graph and all following ones, the positive scale values indicate female overrepresentation and the negative scale values indicate male overrepresentation. We have used the following abbreviations for the ILO occupations: PF = professional; MA = manager; CL = clerical; SA = sales; SR = service; PR = production.

three-way association between occupation, sex, and country.¹⁶ This model can be represented as follows:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{(Z_i v_{jk})}, \quad (6)$$

¹⁶ It is no easy task to choose among scalar indices when the multiplicative shift model fails to fit. In this context, one must either (1) distort the data at the point of estimating the model, or (2) apply a model that fits perfectly and then summarize the many parameters of that model in a single scalar index (thereby introducing "distortions" of a different sort). Although some information will necessarily be lost with either approach, we advocate the latter one because the resulting index (i.e., A) is sensitive to all departures from perfect integration rather than merely those which emerge under a particular representation of the common segregation profile.

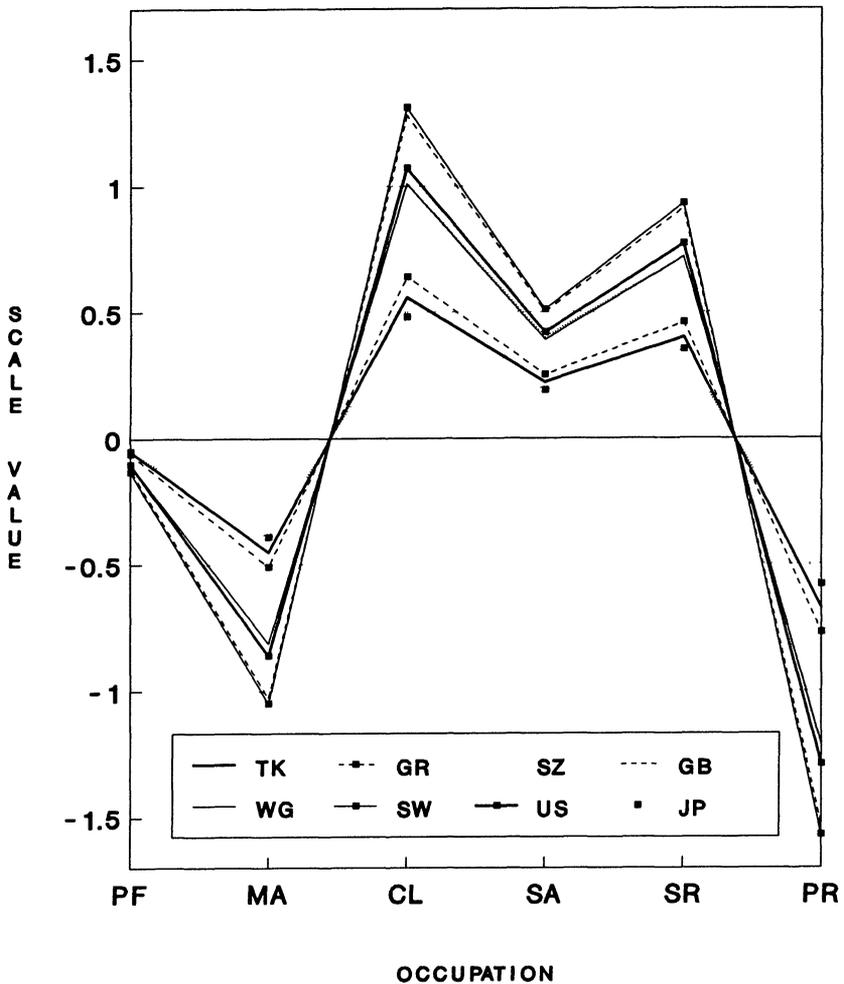


FIG. 3.—Country-specific levels of sex segregation. The parameter estimates are taken from model 3 (table 1). We have used the following abbreviations for the ILO occupations: PF = professional; MA = manager; CL = clerical; SA = sales; SR = service; PR = production. We have also abbreviated the names of countries: TK = Turkey; GR = Greece; SZ = Switzerland; GB = Great Britain; WG = West Germany; SW = Sweden; US = United States; JP = Japan.

where each of the letters and subscripts is defined as before.¹⁷ The column effects under this saturated model can be used to calculate the expected occupation-specific deviation from perfect integration:

$$\begin{aligned} A_k &= \exp\left(1/J \times \sum_{j=1}^J v_{jk}^2\right)^{1/2} \\ &= \exp\left(1/J \times \sum_{j=1}^J \left\{ \ln(F_{jk}/M_{jk}) - \left[1/J \times \sum_{j=1}^J \ln(F_{jk}/M_{jk})\right] \right\}^2\right)^{1/2}. \end{aligned} \quad (7)$$

This new index of association is closely related to ϕ_k .¹⁸ Indeed, when our multiplicative shift model fits perfectly, ϕ_k is completely governed by the size of A . If, for example, A_s and A_t denote the value of A in any two (arbitrarily chosen) countries, then the following result holds:

$$\ln(A_s) = \phi_s/\phi_t \times \ln(A_t). \quad (8)$$

We can therefore conclude that $\ln(A_s)/\ln(A_t)$ will equal ϕ_s/ϕ_t whenever equation (3) correctly characterizes the observed data.

Unlike the size-standardized index of dissimilarity, our new index is invariant under multiplicative transformations of the sex ratio. When F_j is multiplied by the factor c for all j , A can be reexpressed as follows:

$$\begin{aligned} A'_k &= \exp\left(1/J \times \sum_{j=1}^J \left\{ \ln(cF_{jk}/M_{jk}) - \left[1/J \times \sum_{j=1}^J \ln(cF_{jk}/M_{jk})\right] \right\}^2\right)^{1/2} \\ &= \exp\left(1/J \times \sum_{j=1}^J \left\{ \ln(c) + \ln(F_{jk}/M_{jk}) \right. \right. \\ &\quad \left. \left. - \left[1/J \times \sum_{j=1}^J \ln(c) + \ln(F_{jk}/M_{jk})\right] \right\}^2\right)^{1/2} \\ &= \exp\left(1/J \times \sum_{j=1}^J \left\{ \ln(F_{jk}/M_{jk}) - \left[1/J \times \sum_{j=1}^J \ln(F_{jk}/M_{jk})\right] \right\}^2\right)^{1/2} \\ &= A_k. \end{aligned} \quad (9)$$

¹⁷ The marginal effects for this model are identified as before (see eq. [4]), whereas the column effects are now constrained to sum to zero within each country.

¹⁸ In an earlier draft of this article, we used the absolute value function in defining A (see also Charles 1992). The present definition has the virtue of making the relationship between A_k and ϕ_k more transparent (see eq. [8]).

TABLE 2
 SCALAR MEASURES OF OCCUPATIONAL
 SEX SEGREGATION APPLIED TO
 EIGHT-NATION DATA SET

COUNTRY	SEGREGATION INDEX		
	<i>D</i>	<i>D_s</i>	<i>A</i>
Switzerland	39.9	41.6	3.53
Sweden	41.2	41.9	3.07
Great Britain.....	44.4	41.4	2.77
Turkey.....	40.5	46.2	2.64
Japan	24.1	30.1	2.56
Germany	38.9	34.2	2.41
United States.....	36.6	28.9	2.41
Greece	30.2	27.8	1.90

NOTE.—*D* = index of dissimilarity; *D_s* = size-standardized index of dissimilarity; *A* = global association index under saturated model. The values of *A* are calculated from the saturated model (see eq. [6]).

As shown here, the constant factor “cancels out” under the algebra of natural logarithms, and *A'* thus reduces to *A*. Moreover, *A* is also unchanged when the *j*th occupational margin is multiplied by a constant, because $(cF_{jk}/cM_{jk}) = (F_{jk}/M_{jk})$ for any arbitrary *c*. It follows from this result that *A* can be safely used to compare countries with different occupational distributions.

The values of *D*, *D_s*, and *A* for our eight-nation data are presented in table 2. This table indicates that the rank ordering of countries under our margin-free index differs in nontrivial ways from what prevails under *D* or *D_s*. For example, Switzerland registers the highest segregation level under *A*, while it occupies a middle position under *D*. We also find that Japan and Turkey are outliers under *D* and *D_s* but not under *A*. The conventional segregation scores for Japan are usually regarded as surprisingly low (see Brinton and Ngo 1991); however, the present results indicate that the intrinsic association in Japan is quite strong, with our new index implying that males or females are overrepresented in the average Japanese occupation by a factor of 2.56.

CONSTRUCTING SEX SEGREGATION PROFILES

The test statistics presented in table 1 (see contrast 2) make it clear that any scalar index will conceal cross-national variability in the underlying segregation profile. As indicated earlier, we do have some hypotheses

TABLE 3

OCCUPATION-SPECIFIC SEX SEGREGATION PARAMETERS FOR EIGHT-NATION DATA SET

COUNTRY	COLUMN EFFECT					
	Professional	Manager	Clerical	Sales	Service	Production
Turkey.....	<i>1.28</i>	-.74	<i>1.41</i>	-1.01	-.40	-.55
Greece.....	.61	-.82	.84	-.19	.33	-.77
Sweden.....	.22	-1.29	<i>1.50</i>	-.10	<i>1.16</i>	-1.50
Japan.....	.55	-1.90	.87	-.03	.74	-.23
United States.....	.01	-.53	<i>1.45</i>	.01	.49	-1.43
Switzerland.....	.19	-2.20	.80	.93	<i>1.39</i>	-1.12
Great Britain.....	-.22	-1.02	<i>1.27</i>	.58	.92	-1.52
Germany.....	.10	-1.14	.91	.69	.68	-1.24

NOTE.—The parameter estimates are taken from the saturated model (see eq. [6]). The positive estimates indicate female overrepresentation, and the negative estimates indicate male overrepresentation. Estimates that are large and positive are shown in italic.

about the contours of these cross-national differences, but for our present purposes it is useful to proceed inductively by grouping the unconstrained column effects into distinct segregation profiles. The results from this exercise are presented in table 3 and figure 4.

The column effects in table 3 provide limited evidence of cross-national parallelism in occupational sex typing. In all eight countries, we find that women are concentrated in the vast “middle class” of clerical work, while men consistently dominate managerial and production occupations. It would be a mistake, however, to gloss over the rather substantial cross-national differences in the *degree* of sex segregation within these categories. For example, Japanese males are overrepresented in production work by a factor of only 1.26, whereas the corresponding sex ratios for Great Britain and Sweden are as high as 4.57 and 4.48 ($\exp[.23] = 1.26$; $\exp[1.52] = 4.57$; $\exp[1.50] = 4.48$).¹⁹ We would further note that the American managerial sector is highly integrated by current international standards; indeed, American males are only 1.70 times more likely than their female counterparts to be managers, whereas the corresponding sex ratios for Sweden, Japan, and Switzerland are two to five times stronger ($\exp[0.53] = 1.70$; $\exp[1.29] = 3.63$; $\exp[1.90] = 6.69$; $\exp[2.20] = 9.03$).²⁰ These results are clearly consistent with some of the

¹⁹ We have reversed the sign of the coefficients in table 3, since we are now reporting the male-to-female sex ratio.

²⁰ The contrast between the Swedish and American sex ratios may reflect the different ways in which these countries have accommodated demands for reduced gender stratification. As argued elsewhere, legalistic guarantees of equal opportunity (e.g., “affirmative action” policies) have historically played a lesser role in Swedish-style corporatist systems, since interest groups in such systems tend to be co-opted with more

country-specific institutional characteristics that we briefly discussed in our introductory comments.

The sex ratios for professional, sales, and service occupations provide additional evidence that the segregation systems of industrial societies have not yet converged to a common pattern. In each of these occupations, we find cross-national inconsistencies in the direction of occupational sex typing; it is these types of "sign shifts" in the column effects that account for the poor fit of our baseline specification (see table 1, model 3). The most extreme examples of such shifts can be found in the segregation profiles for Turkey and Greece. As shown in figure 4, sales occupations are male dominated under profile A (see esp. Turkey), while the same occupations are gender neutral under profile B and female dominated under profile C. Obversely, we find that the professional sector is female dominated under profile A but gender neutral (or more nearly so) under profiles B and C.²¹ Although results of this general sort could always be attributed to cross-national differences in the composition of the major occupations (see below for details), we suspect that true sociocultural forces can account for some of the sign shifts. It is hardly surprising, for example, that Turkish women have been driven away from sales and service work, since traditional Islamic law strictly regulates the role of women in commercial transactions (see Arat 1989).

The foregoing graphs thus indicate that Turkey and Greece have a top-heavy pattern of segregation. We can further distinguish between two types of bottom-heavy profiles: the first one has a bimodal cast (see profile B), whereas the second one is less jagged in the interior regions (see profile C). The resulting typology can be usefully formalized by estimating a model that fits a single segregation curve for each profile. As shown in table 1, the latter model fits quite well; in fact, our profile-specific specification correctly allocates 98.7% of the respondents (see table 1, model 4), and it accounts for approximately 84% of the total cross-national variability (see table 1, contrast 3). The remaining variability is partly attributable to differences in the degree of sex segregation within each profile.

tangible concessions (e.g., state-financed child care and guaranteed pregnancy leaves). The pluralistic and legalistic political structure of the United States has been translated into affirmative action programs rather than state-financed concessions of the Swedish sort (see Lovenduski 1986; Silén 1988; Gelb 1989; Charles 1990, 1992).

²¹ It could well be argued that the sex segregation regime of Japan is more properly classified under profile A. In this regard, we should emphasize that our profiles obviously cannot be viewed as definitive, but rather are merely illustrative representations of the types of log-multiplicative analyses that might be carried out.

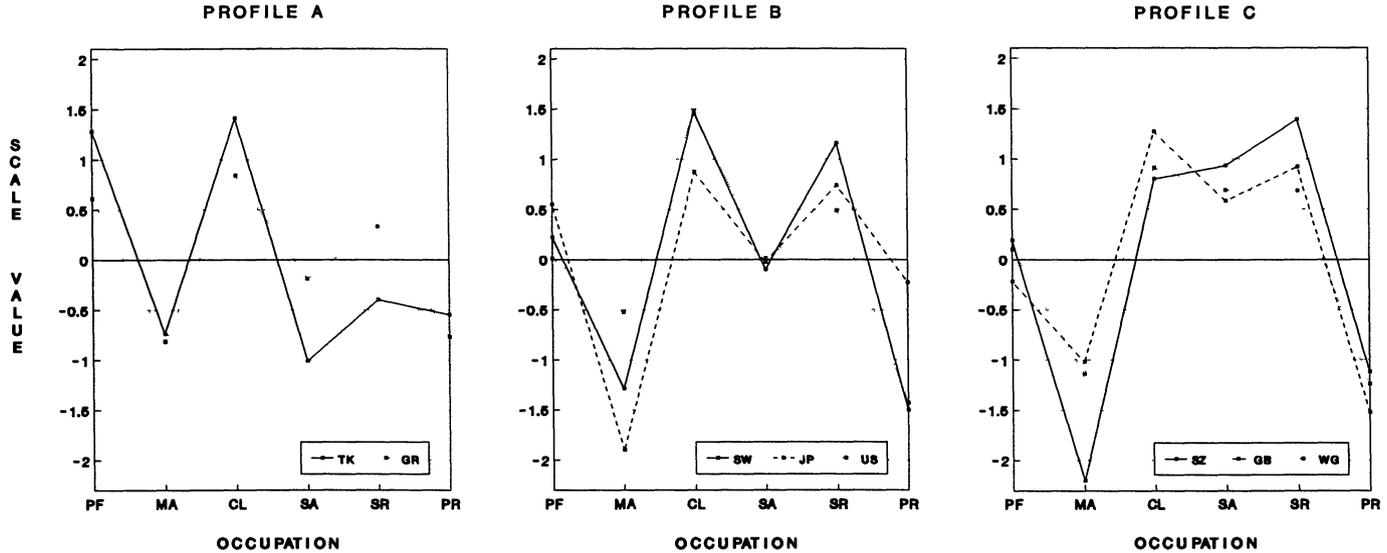


FIG. 4.—Three profiles of occupational sex segregation in eight-nation data set. The parameter estimates are taken from the saturated model (see eq. [6]). We have used the following abbreviations for the ILO occupations: PF = professional; MA = manager; CL = clerical; SA = sales; SR = service; PR = production. We have also abbreviated the names of countries: TK = Turkey; GR = Greece; SZ = Switzerland; GB = Great Britain; WG = West Germany; SW = Sweden; US = United States; JP = Japan.

MODELS FOR DISAGGREGATED DATA

Up to now, our analyses have been based exclusively on the aggregated ILO classifications, but clearly the same types of models and methods could be readily applied to disaggregated data. The purpose of the present section is to introduce a series of multiplicative models that provide new insights into the effects of aggregation and disaggregation on segregation statistics. We will address the following questions in turn:

1. How much of the total sex-by-occupation association is lost by aggregating the data into six ILO categories? Can we capture the most important cross-national differences in sex segregation with highly aggregated data?
2. Are the aggregated ILO profiles (see fig. 4) distorted by cross-national differences in the mixture of detailed occupations found within each major category? Does a cross-nationally common curve emerge when these compositional biases are purged from the data?
3. Can we characterize the detailed segregation profiles with a multiplicative shift parameter? Does the rank-ordering of countries under this shift parameter mimic the corresponding rank-ordering for aggregate data (see table 2)?

As we noted earlier, this set of supplementary analyses will have to be carried out with our two-nation sample (see app. B), since the ILO classifications are only published in highly aggregated form. The results presented in this section should therefore be seen as largely illustrative in intent.

We will begin by asking whether the aggregated ILO classification conceals a substantial amount of sex-by-occupation association. This question can be addressed by fitting the following model for each country:

$$m_{ij} = \alpha\beta_i\gamma_j e^{(Z_i v_c^*)}, \tag{10}$$

where

$$\begin{array}{ll} c = 1 & \text{if } 1 \leq j \leq 13, \\ c = 2 & \text{if } 14 \leq j \leq 15, \\ c = 3 & \text{if } 16 \leq j \leq 22, \\ c = 4 & \text{if } 23 \leq j \leq 25, \\ c = 5 & \text{if } 26 \leq j \leq 31, \\ c = 6 & \text{if } 32 \leq j \leq 45. \end{array}$$

The above set of “side constraints” on v_c^* forces the column effects to be equal within each of the six major ILO categories.²² We have not fitted a corresponding set of column effects for detailed occupations; conse-

²² This model can be identified by fixing β_1 and γ_1 at “1” and by requiring the six column effects (v_c^*) to sum to “0.”

TABLE 4
COMPONENTS OF TOTAL ASSOCIATION IN DISAGGREGATED DATA
FROM UNITED STATES AND JAPAN

Model	L^2	df	L_h^2/L_i^2	Δ
United States:				
Independence ($O + S$)	6,902,112	44	100.0	25.6
Major occupation effects ($O + S + G$)...	3,219,930	39	46.7	13.1
Japan:				
Independence ($O + S$)	2,481,293	44	100.0	20.1
Major occupation effects ($O + S + G$)...	1,663,678	39	67.0	12.8

NOTE.— O = detailed occupation; S = sex; G = major occupation column effects.

quently, the model implies that the segregation regime is homogeneous across occupations after the ILO effects are parsed out, with the associated test statistic thus indexing the extent to which the data violate this implication (see Goodman 1981*b*; Bishop, Fienberg, and Holland 1975, pp. 126–30; Featherman and Hauser 1978, pp. 180–84; Allison 1980). As shown in table 4, 46.7% of the association in the American table is lost by aggregating the data into six ILO categories, while a full 67.0% of the Japanese association is lost under the same aggregation. In interpreting these results, some scholars might prefer to emphasize that a relatively large component of association can be explained with only 5 df , whereas others might point out that most of the sex-by-occupation association is generated within the six ILO categories. The latter type of description tends to be preferred by contemporary segregation researchers; that is, the prevailing view seems to be that the standard occupational groupings (e.g., the ILO categories) are unacceptably heterogeneous and that researchers should therefore attempt to ratchet segregation analyses down to the lowest possible level (see, e.g., Cain 1984; Reskin and Roos 1987, p. 11; Sokoloff 1987, p. 62).²³ We see nothing in the results of table 4 that is inconsistent with such an interpretation.

It is quite another matter to ask whether aggregated data will typically misinform us about the structure of cross-national variability in sex segregation. The sources of this variability can be specified with the following model:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{(Z_i v_j + Z_i v_{ik}^*)}, \quad (11)$$

²³ This viewpoint is typically informed by the pathbreaking work of Bielby and Baron (1984, 1986).

TABLE 5
 LOG-LINEAR AND LOG-MULTIPLICATIVE ASSOCIATION MODELS
 APPLIED TO TWO-NATION DATA SET

Model	L^2	df	L^2_h/L^2_i	Δ
Constant sex segregation ($O \times N + S \times N + C$).....	730,843	44	100.0	4.9
Additive shift in means ($O \times N + S \times N + C + G \times N$).....	220,520	39	30.2	2.0
Hybrid model ($O \times N + S \times N + C + A \times N + G \times N$)...	219,457	38	30.0	2.0

NOTE.— O = detailed occupation; N = country; S = sex; C = detailed occupation column effects; A = global row-by-column association parameter; G = major occupation column effects.

where v_j refers to the unconstrained column effects and v_c^* refers to the ILO column effects that are generated by the six sets of equality constraints implied by equation (10).²⁴ Under the specification indicated here, the only outlet for cross-national variability is the shift effects defined at the level of major categories, since the scale values for the detailed occupations are constrained to be the same across countries. It turns out that this type of constraint does not lead us too far astray; in fact, the model represented by equation (11) misclassifies only 2% of the respondents, and it accounts for nearly 70% of the total cross-national variability (see table 5, “Additive shift in means”). We can thus conclude that the most important institutional differences in sex aggregation are expressed at the level of major occupational groupings. Although the ILO categories clearly conceal a substantial amount of sex-by-occupation association (see table 4), it would appear that this residual association takes on a relatively similar form in each country.

The latter result does not give us full license to proceed with an aggregate analysis. After all, the occupational composition of the ILO categories may well differ across countries, and it is therefore possible that some of the cross-national variability observed in figure 4 is artifactual (see Sokoloff [1987, pp. 63–66] for a related argument). We can eliminate this form of “compositional bias” by fitting the following model:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{(Z_i v_{jk} + Z_i v_{ck}^*)}, \tag{12}$$

²⁴ The identifying restrictions deployed here are directly analogous to those represented by eq. (4). That is, the set of six ILO column effects for the second country (v_{c2}^*) are constrained to sum to zero, and the pooled set of 45 microlevel column effects (v_j) are likewise constrained to sum to zero. The ILO column effects for the first country (v_{c1}^*) are all fixed at one.

where v_j and v_c^* are defined as before.²⁵ The distinctive feature of this model is that it parameterizes the structure of segregation at multiple levels, thereby making it possible to estimate the net residue of segregation at the aggregate level after first purging the data of lower-order compositional effects (see Stier and Grusky [1990] for a related model; also, for a pathbreaking discussion of purging, see Clogg and Eliason [1988] and Clogg, Shockey, and Eliason [1990]). We have graphed the relevant results from this model in figure 5, with the dotted lines denoting the purged ILO column effects, and the solid lines denoting the unpurged effects calculated from the aggregated data.²⁶ The overall picture that emerges from this figure is only partially reassuring; to be sure, some of the original cross-national differences persist in unchanged form (i.e., PF and PR), but the remaining ones are either strengthened or weakened. While this finding indicates that the contours of our segregation profiles (see fig. 4) may be somewhat distorted, it is worth noting that the purged column effects in figure 5 continue to vary across the two nations (see esp. MA, SA, and PR). We must therefore reject our provisional hypothesis that such biases were concealing a cross-nationally common profile.

In some circumstances, analysts of sex segregation may also wish to characterize the structure of the detailed segregation curve, and their attention will thus turn to some form of multiplicative shift model. As may be recalled, we were forced to reject this highly restrictive specification for the aggregated ILO data (see table 1, contrast 2), but it is still possible that the segregation profile for detailed occupations will take on a more structured form. This possibility can be addressed by fitting the following model:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{[\phi_k(Z_i v_j) + Z_i v_{ck}^*]}, \quad (13)$$

where ϕ_k is a multiplicative shift parameter governing the relative height of the peaks and valleys in the detailed segregation profile.²⁷ As indicated

²⁵ The model presented here will always fit the data perfectly. The microlevel column effects (v_{jk}) were identified by being forced to sum to zero within the six major occupational categories of each country, whereas the ILO column effects (v_{ck}^*) were identified by being forced to sum to zero within each country. When these restrictions are imposed, the microlevel column effects account for 39 *df* in each country (i.e., 45 - 6 = 39), and the ILO column effects account for the remaining 5 *df* in each country (i.e., 6 - 1 = 5).

²⁶ The unpurged column effects in fig. 5 differ from the corresponding effects in fig. 4, since the two sets of estimates are based on different data.

²⁷ The shift effects (ϕ_k) for this model were identified by fixing ϕ_1 at one, and the column effects (v_j and v_{ck}^*) were identified by imposing the restrictions specified in n. 24. For didactic purposes, we have presented two sets of means (v_{c1}^* and v_{c2}^*) in fig. 6, but it should be kept in mind that one of these sets is implied by the microlevel column effects and does not, therefore, convey any additional information.

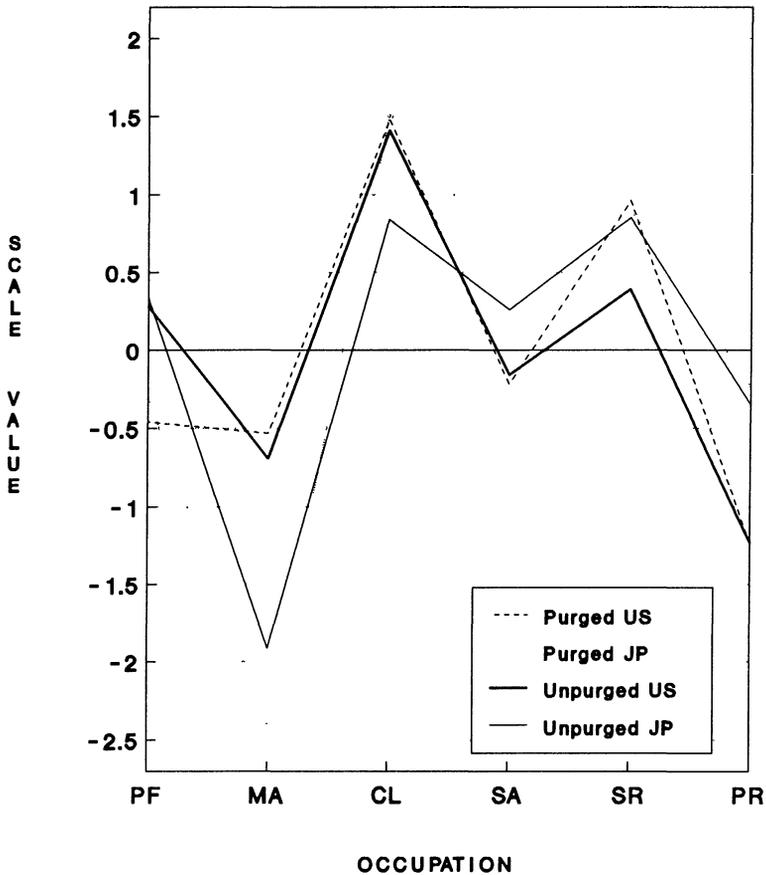


FIG. 5.—Purged and unpurged means for two-nation data. The purged scale values are drawn from the saturated model represented in eq. (12). We have used the following abbreviations for the ILO occupations: PF = professional; MA = manager; CL = clerical; SA = sales; SR = service; PR = production. We have also used abbreviated names for Japan (JP) and the United States (US).

in figure 6, this new model permits the purged ILO means (v_j^*) to freely vary across countries, but it constrains the lower-order column effects (v_j) to be a multiplicative function of ϕ_k (see table 5 for the relevant fit statistics). In the present case, the estimated ratio of ϕ_1 to ϕ_2 equals 1.047, and we can thereby conclude that the dispersion around the ILO means is approximately 4.7% greater in Japan than in the United States. It follows that Japan has a higher segregation index at two levels of analysis; that is, not only are the Japanese ILO categories more segre-

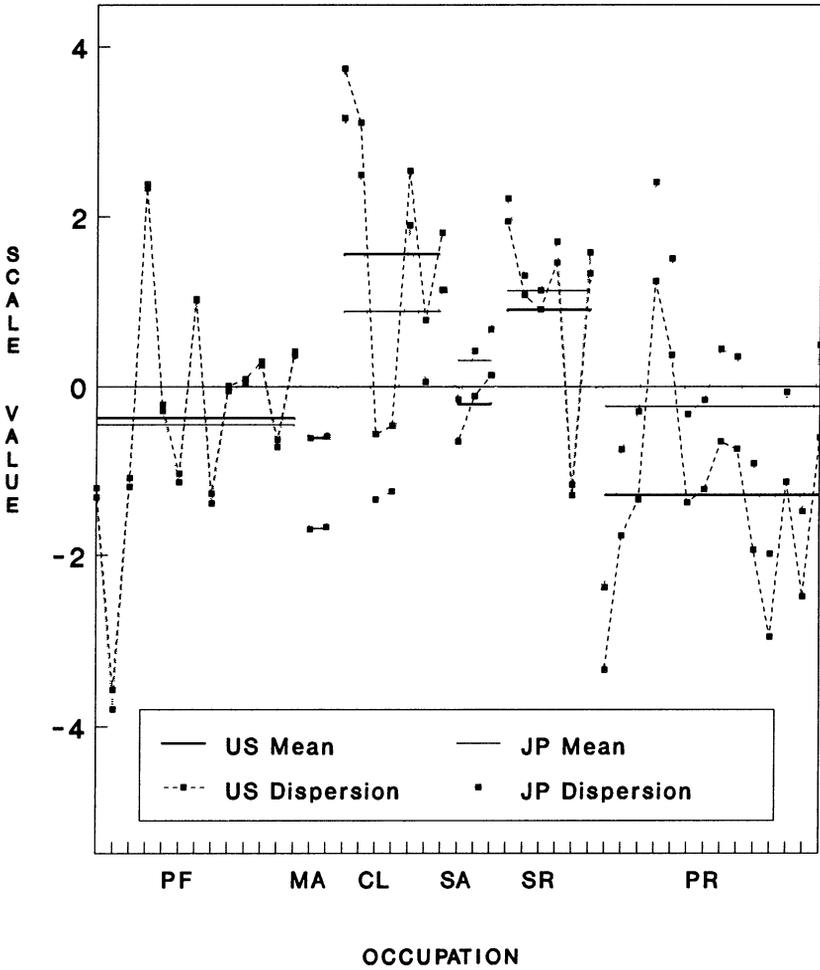


FIG. 6.—Hybrid multiplicative model for two-nation data. The parameter estimates are taken from the hybrid model (table 5). We have used the following abbreviations for the ILO occupations: PF = professional; MA = manager; CL = clerical; SA = sales; SR = service; PR = production. We have also used abbreviated names for Japan (JP) and the United States (US).

gated than the corresponding American categories, but so too are the detailed occupations within these categories.²⁸

THE HIERARCHICAL STRUCTURE OF SEX SEGREGATION

The models that we have presented up to now all treat the occupational structure as a simple nominal variable. Although most of the standard segregation indices (e.g., D , D_s) also rest on a nominal level of measurement, there has been a recent resurgence of work based on alternative indices that require some form of ordinal or continuous scaling (see Brinton and Ngo 1991; Stafford and Fossett 1989; Fossett, Galle, and Kelly 1986; White 1983). This work has been motivated, at least in part, by the so-called checkerboard problem that White (1983) first discussed in the context of residential segregation research. As was noted by White (1983, pp. 1010–11), the value of D will be unaffected by the spatial arrangement of the underlying census tracts (i.e., their “checkerboard layout”), since the relevant calculations for D are based on the internal composition of the tracts rather than the distances between them (see also Duncan and Duncan 1955, p. 215; Taeuber and Taeuber 1965, p. 205). The implication, of course, is that D will fail to register “tract-level” desegregation; for example, if a large ghetto were broken up by scattering the constituent tracts throughout the city, the value of D would necessarily remain unchanged.²⁹

The same type of problem arises in a simpler (one-dimensional) form when occupational data are analyzed. In the latter context, one might wish to calculate the “social distance” between the male and female distributions, with the relevant metric typically being some form of prestige or socioeconomic scale. It is clearly inappropriate to use D for such purposes, since it was designed to measure “nominal differentiation rather than inequality” (Fossett et al. 1986, p. 423). Indeed, just as D cannot detect the residential desegregation that occurs when an all-black tract is moved to a neighborhood formerly dominated by whites, so too it cannot detect the sex desegregation that occurs when an all-female occupation is moved to a “socioeconomic region” formerly dominated by males (see also Fossett and South 1983; Stafford and Fossett 1989, p.

²⁸ If the multiplicative shift effects are allowed to be category specific, we find that in two cases (i.e., clerical and production occupations) the segregation index is actually larger in the United States than in Japan. This elaborated model fits relatively well ($L^2 = 193,005$), but it still accounts for only 12.5% of the total variability in the detailed segregation profile ($1 - [193,005/220,520] = .125$). We are thus well advised to treat the estimates from our multiplicative shift models with some caution.

²⁹ The obvious irony here is that the checkerboard problem can be solved (in ad hoc fashion) by resorting to *less* detailed levels of measurement.

179; Brinton and Ngo 1991). This deficiency has motivated some segregation scholars to define and deploy alternative indices that take into account the location of census tracts, school districts, or occupations in physical or social space. It should come as no surprise that these revised indices are often direct modifications of D (e.g., Brinton and Ngo 1991).

Unless the indices so proposed are functions of the relevant cross-product ratios, they will again be margin-dependent and therefore flawed for comparative purposes. We can secure a margin-free measure by fitting a scaled association model of the following kind:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{\psi_k(Z_i T_{jk})}, \quad (14)$$

where ψ_k and T_{jk} are direct analogues to ϕ_k and v_{jk} , and the remaining parameters retain their original meaning. This model differs from our baseline specification in equation (3) because T_{jk} refers to a priori values rather than freely estimated ones.³⁰ We thus end up with a hybrid specification that stands somewhere between the association models of Haberman (1974) and those of Hout (1984). As indicated in equation (14), the row categories in our data are scaled with the standard unit scores of a linear-by-linear interaction model (see Haberman 1974; Duncan 1979; Goodman 1979a), whereas the column categories are scaled with external scores of the kind deployed by Hout (1984, 1988), Hauser (1984), and others (Hout and Jackson 1986; Szelényi 1988).

We have illustrated this simple approach by estimating ψ_k conditional on the prestige scores from SIOPS (Standard International Occupational Prestige Scale).³¹ Under our specification, T_{j1} equals T_{j2} for all j ; this constraint holds because we have scaled the occupations in each country with the same prestige scores. The fit statistics in table 6 indicate that our model misallocates 23.7% of the respondents and accounts for only 1.2% of the total sex-by-occupation association (see the scaled association model). It would thus appear that SIOPS cannot adequately account for the underlying structure of sex segregation; if anything, the graphs in figure 7 suggest that the column effects take on a curvilinear form, with the inflection point in both countries occurring between 45 and 50 prestige points. This poor performance is of course consistent with some of our prior results. The graphs in figure 5, for example, suggested that the ILO

³⁰ The multiplicative shift parameter indexes the overall sex-by-occupation association conditional on the scores so assigned.

³¹ In almost all cases, the occupations in our 45-category classification correspond to a "minor group" in ISCO, and we could therefore directly apply the published ISCO version of the standard scale (see Treiman 1977, app. A). The remaining scores were estimated by averaging across the SIOPS values for all of the "unit occupations" contained within a given category (see app. B below for a listing of the final scores).

TABLE 6
 MODELS OF SCALED ASSOCIATION FOR TWO-NATION DATA

Model	L^2	df	L_h^2/L_i^2	Δ
Conditional independence ($O \times N + S \times N$)	9,383,405	88	100.0	23.7
Scaled association ($O \times N + S \times N + P \times N$)	9,274,708	86	98.8	23.7
Purged scaled association ($O \times N + S \times N + P \times G \times N$)	3,667,937	66	39.1	10.1

NOTE.— O = detailed occupation; N = country; S = sex; P = product term (row scores \times column scores); G = major occupation.

column effects do not follow a simple prestige gradient, whereas our additional tests in table 4 indicated that nearly half of the association in the disaggregated tables is generated at the ILO level. These two findings imply that a simple association model was doomed from the start.

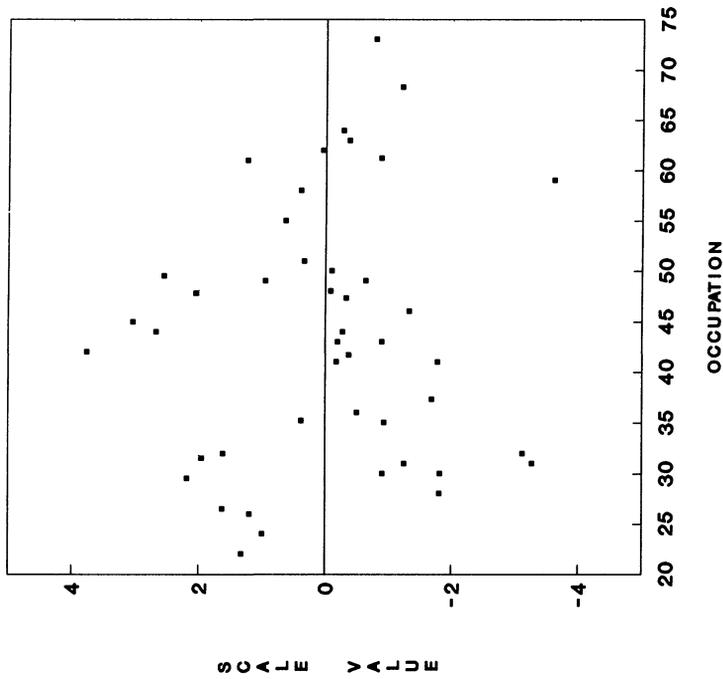
The expected prestige gradient may nonetheless emerge after the ILO effects are purged from the data. It is commonly argued, for example, that aggregate analyses of sex segregation are misleading because males tend to secure the most desirable occupations within each of the major categories conventionally deployed. This expectation can be tested, albeit only partially, by fitting a model of the following kind:

$$m_{ijk} = \alpha_k \beta_{ik} \gamma_{jk} e^{[\Psi_{ck}^*(Z_i T_j) + Z_i \nu_{ck}^*]}, \tag{15}$$

where ν_{ck}^* refers to the ILO column effects in the k th country (see eq. [10] for the relevant equality constraints), Ψ_{ck}^* refers to the corresponding country-specific association parameters estimated within each of the six ILO major categories, and the remaining letters and subscripts are defined as before.³² As might be expected, this revised specification fits relatively well; the results in table 6 reveal that only 39.1% of the association remains unexplained when ν_{ck}^* and Ψ_{ck}^* are included as additional terms. While the fit statistics are much improved under this specification, we still find that the resulting prestige gradient does not take on the conventionally expected shape (see table 7). In fact, males enjoy an advantage over females in only three ILO categories (professional, sales, and service), whereas the reverse association prevails in the two re-

³² This model can be estimated by applying our original association model (see eq. [14]) to the six subtables formed by disaggregating across the ILO categories. If the data are analyzed as a single array (as represented in eq. [15]), then the ILO column effects (ν_{ck}^*) can be identified by being forced to sum to zero within each country.

UNITED STATES



JAPAN

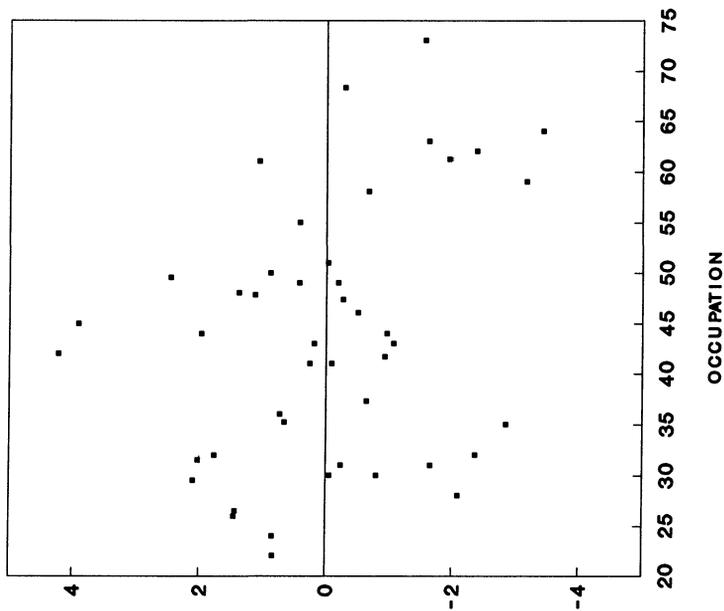


Fig. 7.—Scatter plots of column effects by prestige for the United States and Japan. The scale values are taken from the saturated model. The horizontal axis indexes the SIOPS scores for the detailed occupations in our 45-category classification.

TABLE 7
PARAMETERS OF SCALED ASSOCIATION FOR
MAJOR OCCUPATIONAL CATEGORIES

OCCUPATION	ASSOCIATION INDEX	
	United States	Japan
Professional.....	-.113	-.118
Clerical.....	.217	.112
Sales.....	-.044	-.040
Service.....	-.213	-.184
Production.....	.010	.021

NOTE.—The parameter estimates are taken from model 3, table 6. The positive estimates indicate female overrepresentation, and the negative estimates indicate male overrepresentation. We have omitted the managerial category because it contains only two occupations.

maining cases (clerical and production). If our SIOPS scores were a perfect measure of the general desirability of jobs (see Goldthorpe and Hope 1974), these results would be partially inconsistent with the hypothesis that “occupational composition [is] the result of a matching process in which the top-ranked workers get the most attractive jobs” (Reskin and Roos 1990, p. 307; see also Strober 1984).³³

CONCLUSIONS

We began this article by noting that the conventional segregation indices are dependent on the marginal distributions in a sex-by-occupation array (see fig. 1). It was this deficiency that motivated Gibbs (1965), Gross (1968), and other segregation scholars to modify *D* by standardizing for “differences among occupational categories with regard to their share of the labor force” (Gibbs 1965, p. 163). The resulting size-standardized index (*D_s*) has now become the measure of choice among contemporary scholars who seek to compare segregation regimes across time or space (e.g., Brinton and Ngo 1991, 1993; Presser and Kishor 1991; Williams 1979; Jacobs 1989a, 1989b; Jacobs and Lim 1992). It should be empha-

³³ It should be emphasized, of course, that a queuing hypothesis of this sort cannot be adequately addressed with the present data. After all, SIOPS is clearly not an exhaustive measure of job desirability (see Jencks et al. 1988), nor is our 45-category classification sufficiently fine grained to reveal the structure of male advantage at the most detailed level of analysis (see Bielby and Baron 1984). We would further note that a queuing model cannot be tested in any convincing way without controlling for heterogeneity in worker qualifications (e.g., education, work experience, and training).

sized that the transition to D_s occurred without great fanfare; indeed, D_s became popular well after the methodological debates of the 1960s had run their course, and it was therefore shielded from the rigorous vetting that earlier indices had undergone (see Peach 1975). This is not to say that the standardization proposed by Gibbs (1965) failed in its stated objective to eliminate the distorting effects of the occupational structure. The size-standardized index does indeed live up to its billing; however, the cost of standardizing in this fashion is that D_s is no longer scale invariant (see James and Taeuber 1985), and hence researchers replacing D with D_s are merely exchanging one form of marginal dependence for another.

If we wish to construct a margin-free index, we have no choice but to resort to measures that are functions of cross-product ratios. The centerpiece of our approach has been a simple association model that permits the peaks and valleys of the segregation profile to be compressed or expanded in accord with a scalar shift effect. We have thus rejected the conventional practice of assuming that a scalar index is adequate to the task; instead, we have argued that this assumption should be embedded in a testable model, with the viability of a scalar approach resting on the usual criteria of model fit. As it turns out, only one-third of the total cross-national variability can be explained with a multiplicative shift effect, while the remaining variability must be attributed to heterogeneity in the segregation curves themselves (see table 1, contrast 2). Although the latter form of heterogeneity cannot be fully captured with a scalar index, we have nonetheless defined a margin-free measure (A) that is based on the particular segregation profile prevailing within each country. We have also used elaborated versions of our general model to examine the hierarchical structure of segregation, to identify the dominant "segregation profiles" in industrial countries, and to parse out the net residue of segregation at multiple levels of analysis.

This modeling framework may prove to be useful in future analyses of occupational sex segregation. At the same time, the conventional indices have so far shown remarkable staying power, and it would thus be presumptuous of us to suppose that the forces of inertia and entrenched interest will be quickly or easily overcome. Among the various arguments that are likely to be raised against our analyses, the following are perhaps the most obvious ones:

Issues of aggregation.—By the usual standards of sex segregation research, our occupational categories are highly aggregated, and some skeptics may therefore find our results to be unconvincing or even misleading. In this regard, we would emphasize that our analyses were completed at the aggregate level for reasons of convenience and data avail-

ability, and not because of any intrinsic limitations of our modeling framework.³⁴ We would nonetheless contend that patterns of sex segregation at the major occupational level are of considerable interest because they signify correspondingly major differences in socioeconomic rewards and conditions. In our continuing efforts to ferret out segregation at the most detailed level possible, we ought not to forget that much of the occupational heterogeneity in life chances, work conditions, and consumption practices is likely located at the level of major categories. While the analytic returns to disaggregating occupations may therefore be limited, this is not to gainsay the equally important point that research carried out at the major occupational level is subject to compositional biases that can and should be purged by the methods that we introduced above (see eq. [12]).

Conceptualizing sex segregation.—It is also important to address the fallback argument that log-multiplicative measures are not sufficiently faithful to traditional conceptualizations of sex segregation. If sex segregation is defined to be whatever D or D_s measure, then of course such arguments hold in a nominal sense. We would suggest, however, that many researchers *have* adopted an implicit conceptualization of segregation that is distinct from these common operationalizations (e.g., Williams 1979; Blau and Hendricks 1979; England 1981; Bridges 1982; Handl 1984; Tienda and Ortiz 1987; Beller 1984; Bianchi and Rytina 1986; Abrahamson and Sigelman 1987; Jacobs 1989*a*, 1989*b*; Jacobs and Lim 1992; Presser and Kishor 1991; Brinton and Ngo 1991, 1993). Indeed, in treating column effects as the fundamental parameters of sex segregation, we have merely operationalized the long-standing assumption that such parameters are properly independent of both the occupational structure and the rate of female labor force participation. The development of D_s was seemingly motivated by a similar objective, yet it failed to fully realize the implicit conceptualization that underlies it.³⁵

Issues of endogeneity.—In making the prior point, we do not mean to suggest that traditional conceptualizations of segregation are necessarily desirable, nor that any clear consensus on conceptual matters has now been reached. However, we would argue that scholars who reject margin-free measures of sex segregation should not likewise reject all forms

³⁴ If a cross-classification is disaggregated to the point of being extremely sparse, then asymptotic results are less safely presumed and sampling zeroes may frequently appear. Although some of the models introduced here cannot be directly estimated when sampling zeroes are present, the methods for analyzing such arrays are relatively well developed (see, e.g., Clogg and Eliason 1987).

³⁵ While D also confounds marginal and interaction effects, its great virtue is that it does so in a way that has a pleasing interpretation. The same cannot be said for D_s , and we would therefore recommend abandoning the latter measure altogether.

of multiplicative modeling, since many competing representations of segregation can be elegantly operationalized in the context of simple log-linear or log-multiplicative specifications. This includes, for example, parameterizations in which the rate of female labor force participation is not exogenously determined, but rather is generated by the conjunction of occupational marginal effects and occupation-specific gender ratios. The latter formulation is hardly radical or innovative; it underlies, in fact, all arguments to the effect that the secular trend in female labor force participation has been generated by the exogenous growth of female-typed occupations (see, e.g., Oppenheimer 1970). The following model is suggested by such arguments:

$$m_{ij} = \alpha \gamma_j \delta_{ij}, \quad (16)$$

where $\gamma_1 = 1$ and $\delta_{11} = \delta_{12} = \dots = \delta_{1J} = 1$. The segregation parameters under this specification (i.e., δ_{2j}) are a mixture of the marginal (β_2) and interaction (ν_j) effects estimated under our saturated multiplicative model (see eq. [6]). In a single sex-by-occupation array, these two specifications are perforce equivalent, since both fit the data perfectly. However, when a third dimension is introduced (e.g., time), it becomes possible to determine which of these specifications might be construed as structural and thereby preferred (see Duncan 1975; also, for a relevant application, see Grusky and Hauser [1984]). The model of equation (16) could be elaborated in various other ways, but for our present purposes it should suffice to emphasize that multiplicative models are consistent with a wide range of segregation parameterizations, not all of which are margin free.

The appeal of a multiplicative framework thus rests in large part with its analytic flexibility. We would like to conclude by reviewing, if only briefly, some potentially useful modifications and extensions of our preferred models. Among the more straightforward extensions, we would include (1) models that incorporate additional scalable variables representing further contexts in which segregation processes are nested (e.g., age, period, cohort, industry, and firm size), (2) models that account for cross-context variability in segregation by conditioning on a priori scores that characterize the contexts,³⁶ and (3) models of racial segregation that freely scale both the occupational categories and the racial categories

³⁶ The requisite models here would be similar in structure to those represented in eqq. (14) and (15). However, rather than applying external scores to the occupational categories, we would now be applying such scores to the categories indexing countries. The contextual variables represented in this fashion can be constrained to exert global effects on all occupations or to exert particularized effects that are specific to certain occupations or combinations of occupations (Charles 1992; also, see Grusky and Hauser 1984; Hauser and Grusky 1988).

represented in a race-by-occupation array (e.g., whites, African-Americans, Hispanics, and Native Americans). It would also be useful to revise our preferred models in more ambitious and far-reaching ways by explicitly incorporating individual-level covariates (e.g., education). In doing so, cross-national differences in the endowments or human capital of men and women could be effectively controlled, with the result being a purged version of *A* that indexes the residual variability generated in the labor market itself (see Yamaguchi [1983] for a related model). The latter approach would provide a useful bridge between descriptive measures of sex segregation and standard explanatory models of sex discrimination.

APPENDIX A

TABLE A1
OBSERVED COUNTS IN CROSS-CLASSIFICATION OF SEX
BY OCCUPATION IN EIGHT COUNTRIES

COUNTRY AND SEX	OCCUPATION					
	Professional	Managerial	Clerical	Sales	Service	Production
Turkey:						
Male.....	580,983	148,629	437,380	709,755	833,713	3,675,554
Female.....	244,868	8,310	209,217	30,132	65,342	247,207
Greece:						
Male.....	20,029	5,296	18,311	21,688	18,061	93,755
Female.....	12,440	789	14,310	6,055	8,498	14,744
Switzerland:						
Male.....	290,252	69,673	294,564	110,684	115,560	913,171
Female.....	177,659	3,921	330,847	141,404	235,065	151,017
Great Britain:						
Male.....	2,497,820	1,787,150	1,011,550	571,510	1,004,260	6,796,060
Female.....	1,639,970	524,560	2,933,700	832,620	2,058,480	1,208,680
Germany:						
Male.....	9,528	3,336	8,048	4,236	5,444	30,752
Female.....	6,496	660	12,408	5,232	6,652	5,488
Sweden:						
Male.....	6,006	869	1,089	2,057	1,628	11,407
Female.....	7,183	231	4,675	1,793	4,994	2,453
United States:						
Male.....	69,007	62,813	28,041	53,110	48,578	211,708
Female.....	65,988	34,936	112,601	50,809	74,607	47,986
Japan:						
Male.....	28,710	19,503	44,649	51,381	21,780	147,312
Female.....	23,661	1,386	50,787	23,661	21,780	55,440

NOTE.—We have deflated the population estimates provided by the ILO into sample counts (see Charles 1990).

APPENDIX B

TABLE B1

OBSERVED COUNTS IN DETAILED CROSS-CLASSIFICATION OF SEX
BY OCCUPATION IN UNITED STATES AND JAPAN

OCCUPATION	COUNTRY AND SEX			
	U.S. Males	U.S. Females	Japanese Males	Japanese Females
Professional:				
Researcher and engineer (61.2)...	642,634	157,536	245,666	13,703
Air and ship officer (59.0)	21,202	336	13,285	218
Medical professional (68.3)	122,463	21,672	45,432	13,315
Other medical workers (49.5).....	94,147	704,711	36,709	166,849
Accountant (62.0).....	119,046	73,397	7,917	286
Jurist (73.0)	91,240	24,445	8,026	661
Teacher (61.0).....	286,338	582,908	163,535	184,642
Worker in religion (46.0)	54,687	8,543	18,233	4,288
Author and journalist (58.0)	31,034	27,140	16,930	3,394
Painter and sculptor (51.0)	61,094	50,303	23,481	8,889
Musician and performer (48.0) ...	33,406	18,005	10,645	16,559
Athlete (49.0)	7,644	2,388	4,188	1,345
Other professional (55.0).....	164,548	182,423	33,225	19,639
Manager:				
Government official (64.0).....	48,413	21,634	28,107	355
Business manager (63.0)	1,048,860	425,148	463,655	35,827
Clerical:				
Stenographer and typist (42.0) ...	6,003	151,256	598	16,000
Keypunch operator (45.0)	5,695	69,320	507	9,885
Transport conductor (41.7).....	145,761	58,894	34,524	5,328
Mail clerk (44.0)	119,115	53,483	34,778	5,195
Telephone operator (44.0)	6,482	55,062	7,280	20,494
Machine operator (49.0)	43,163	65,504	10,145	6,048
Other clerical (47.8).....	515,164	2,340,363	806,349	965,553
Sales:				
Sales manager (47.3).....	261,998	110,705	247,146	73,908
Agent and broker (50.0)	170,983	90,689	56,924	53,705
Other sales (35.2).....	620,865	531,291	629,740	474,062
Service:				
Maid and housekeeping (29.5) ...	40,017	208,478	8,298	26,331
Cook and food service (26.0)	312,831	607,134	210,059	346,838
Launderer (22.0).....	17,604	38,756	16,725	14,985
Personal service (32.0).....	33,383	98,422	37,629	84,949
Protective service (35.0).....	585,171	135,701	151,266	3,411
Other service (26.5)	56,372	168,474	37,135	60,555
Production:				
Extractive (32.0).....	66,537	1,726	14,019	517
Metal and plastic worker (37.3)..	537,146	58,543	498,643	101,921
Wood worker (31.0).....	45,608	7,668	73,493	22,621
Textile worker (31.5)	56,147	230,627	94,081	277,981

TABLE B1 (Continued)

OCCUPATION	COUNTRY AND SEX			
	U.S. Males	U.S. Females	Japanese Males	Japanese Females
Shoemaker (24.0)	7,945	12,582	6,291	5,689
Assembler and repairer (43.0)	792,613	189,202	288,013	38,506
Electrical worker (41.0)	296,981	29,446	217,322	109,276
Jeweler (43.0)	4,865	2,350	3,941	1,857
Printer (41.0)	73,689	36,438	57,624	20,540
Painter (30.0)	102,942	9,844	61,258	10,740
Construction worker (31.0)	734,172	16,250	555,789	41,301
Machine operator (30.0)	611,655	144,905	170,537	62,691
Transport equipment (28.0)	564,560	54,343	471,832	22,686
Other production (36.0)	1,177,269	416,955	356,730	289,370

NOTE.—The entries in parentheses are the estimated Standard International Occupational Prestige scores (Treiman 1977).

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