Symmetry and Hierarchy in Social Mobility: A Methodological Analysis of the CASMIN Model of Class Mobility
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Symmetry and hierarchy in social mobility: a methodological analysis of the CASMIN model of class mobility

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ABSTRACT The CASMIN model of class mobility proposed by Erikson and Goldthorpe advances our understanding of cross-national differences in social mobility in a number of important ways, most notably by showing how differences in the association between social origins and destinations reflect consequences of public policies that enhance or restrict opportunities. We respecify the CASMIN model in ways that clarify (*a*) the role of socio-economic differences among classes in mobility processes and (*b*) the extent of cross-national variation. In particular, the problems with the CASMIN model are its application to highly aggregated occupational classes, its suppression of hierarchical or vertical differences among classes, and its asymmetric classification of origin and destination classes. Our alternative specification is based on greater occupational detail, incorporates continuous covariates in linear-by-linear expressions that are analogous to regression models, and imposes symmetry on the association between origins and destinations. We find that the CASMIN model understates the importance of hierarchy relative to sector and inheritance in the determination of mobility patterns generally as well as in cross-national differences. Furthermore, the symmetry of our model facilitates the analysis of structural mobility as a factor that contributes to cross-national differences in overall mobility rates.

INTRODUCTION

The CASMIN project, Comparative Analysis of Social Mobility in Industrial Nations, has produced a wealth of new information on comparative class mobility. Beginning with a common classification scheme more detailed than any attempted for as many countries before,[1] Erikson and Goldthorpe (1987*a*, 1987*b*, 1992) have spelled out new ways of looking at and thinking about social mobility in industrial nations. They find that, throughout Europe, mobility is a multi-dimensional process that depends less on a vertical or hierarchical ordering among social classes and more on sectoral and market differences than sociologists commonly suppose.[2] They show how mobility processes differ from nation to nation and conclude that public policy affects social mobility. However, each nation's experience with mobility policy differs so much from its neighbour's experience that it is difficult to generalize about the effects of such policies.

The key to these conclusions is Erikson and Goldthorpe's core model of the association between social origins and destinations in aggregated inter-generational class mobility tables. We do not address the national variants (Erikson and Goldthorpe, 1987*b*) of the CASMIN model, so we refer to the core model as the CASMIN model. The CASMIN model specifies three hierarchical levels of social stratification, captured in two terms that combine the formal properties of 'crossing parameters' and 'distance parameters' (Goodman, 1972). Mobility within a level is uniformly easier than mobility between levels. The CASMIN model also specifies three levels of occupational inheritance: a baseline level for routine non-manual workers (classes III*a* and III*b*), manual workers (classes V, VI, and VII*a*), and farm labourers (class VII*b*); a second level for professionals (classes I and II) and proprietors (classes IV*a* and IV*b*); and a third level for farmers (class IV*c*). A sectoral effect separates agricultural and non-agricultural workers. Two special 'affinity' parameters account for less-than-expected mobility between the top two classes and the bottom one (I and II vs. VII*b*—a contrast that might better be termed 'disaffinity') and greater-than-expected mobility for a variety of origin–destination combinations. There are four symmetrical combinations of classes included in the second affinity term: mobility between (I,II) and (III*a*,*b*), (I,II) and (IV*a*,*b*), (IV*a*,*b*) and (IV*c*,*d*), and (V,VI) and (VII*a*). There are also two asymmetrical combinations of classes in the second affinity term: from (IV*c*,*d*) to (VII*a*) and from (VII*b*) to (VII*a*).

While the hierarchy, inheritance, and sector parameters in the CASMIN model have a clear theoretical rationale, those for affinity do not. They result from an exploratory analysis and appear to be chosen deliberately to fit the data with a small number of additional parameters. This can be a legitimate procedure (see e.g. Hauser, 1979), inviting future cross-validation, but it leaves the CASMIN model less compelling than one based entirely on theoretical expectations. We have made a modest effort to cross-validate the CASMIN model in an independent sample, and, as we report later in the paper, we find less support for the affinity parameters than for other parts of the CASMIN model.

Despite the weak foundation of the affinity terms, they give the CASMIN model an inherent advantage in comparisons between its fit and the fit of other models. Other parts of our analysis rest on efforts to improve on the fit of the CASMIN model by specifying alternative models that make use of measured variables. The found-in-the-data status of the affinity parameters make this task more difficult than it would be if the CASMIN model contained no terms that were formulated in the process of data analysis.

We use the same data-set as Erikson and Goldthorpe (Erikson *et al.*, 1989) and draw on our own previous work (Hauser 1984*a*, 1984*b*; Hout 1984*a*, 1988) to critique and modify the CASMIN model. The results of our revisions lead us to hone some of Erikson and Goldthorpe's points and reconsider others. Our strategy is to elaborate the coding of the data and to simplify the model. In doing so, we uncover evidence that is not visible with the combination of categories and models that Erikson and Goldthorpe apply. Our results confirm the importance of vertical effects on social mobility, while reinforcing Erikson and Goldthorpe's finding of strong sectoral influence.

The CASMIN data also contain valuable information on structural mobility, i.e. mobility that is forced by factors that are independent of social origins.[3] Structural mobility is a key component of gross mobility differences over time (Hauser *et al.*, 1975; Hout, 1988) as well as cross-nationally (Erikson *et al.*, 1979). Erikson and Goldthorpe give almost no attention to structural mobility. We estimate structural effects on mobility, and we report structural mobility multipliers for each combination of nation and occupation (Sobel *et al.*, 1985).

OCCUPATIONAL CATEGORIES

Heterogeneity in prestige, education, and income

We begin by expanding the classification scheme used by Erikson and Goldthorpe. Problems with data comparability preclude the kind of detailed analysis that is possible when working with a single data-set (e.g. Hout, 1989: 42–50), but we pursue the issue to the limits of the available data. The point of using a more detailed occupational classification is to utilize as much information as possible about the pattern of social mobility in order to improve estimates of the size of hierarchical effects and cross-national differences. Occupational classes that Erikson and Goldthorpe combined prior to analysing the CASMIN data differ widely in socio-economic standing or prestige. The prestige of upper-grade professionals and managers (class I) exceeds that of lower-grade professionals and managers (II) by 12 points on the Hope–Goldthorpe scale (Goldthorpe and Hope, 1974; see Hout, 1989, Table 2.9). Clerical and routine non-manual workers (III*a*) score higher than sales and service workers (III*b*) by 11 points (new calculation). Proprietors (IV*a*) and farmers with employees (IV*c*) score higher than their counterparts without employees (IV*b* and IV*d*) by 3 and 21 points, respectively.[4] Technicians and foremen (V) score 9 points higher than skilled workers (VI). Semi-skilled manual workers score 12 points higher than unskilled workers (both VII*a*). The publicly available data include all of these distinctions except, unfortunately, that between semi-skilled and unskilled workers, and we have reinstated them in our analysis.

The heterogeneity of Hope–Goldthorpe prestige scores among the seven CASMIN classes is displayed in the top panel of Figure 1. Furthermore, there is not a clean break between the two hierarchical divisions recognized among destinations in the CASMIN model. These data are too nearly continuous for discrete classification. That is, the prestige differences among classes raise doubts in our minds as to why upper and lower grade professionals and managers (I and II) should be combined at the top of the hierarchy, unskilled workers (VII*a* and VII*b*) combined at the bottom, and all others combined in the middle. For example, farmers with employees (IV*c*) are closer in prestige to professionals and managers (I and II) than to any of the classes in the middle CASMIN stratum, and farm workers (VII*b*) have virtually the same prestige as sales and service workers (III*b*). Not only are the CASMIN categories internally heterogeneous in prestige, but the vertical hierarchy specified in the CASMIN model does not correspond well with inter-class differences in prestige.

In the two lower panels of Figure 1, we consider two other empirical representations of the vertical dimension of class: education (percentage with more than primary schooling) and income (more than £2500), both observed in the surveys of Ireland and Northern Ireland. Again, in each of these cases there is substantial heterogeneity within categories of the CASMIN

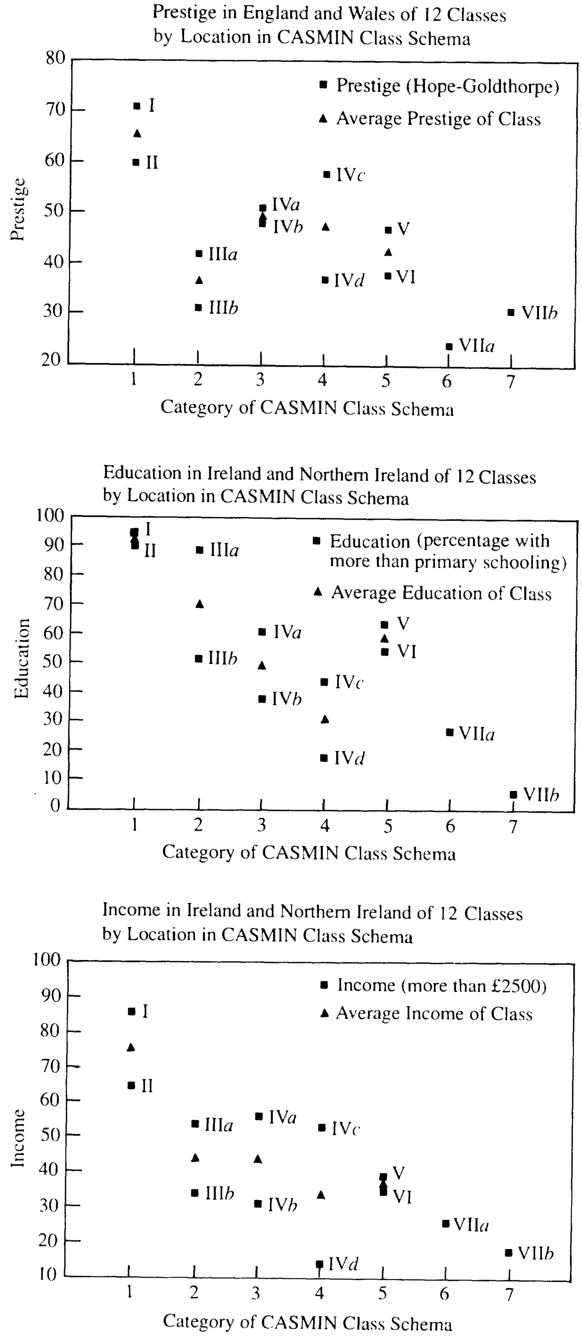

FIGURE 1 *Hierarchy and heterogeneity of 12 occupational classes.*

class schema. Furthermore, there is even less evidence of correspondence between these two criteria and the CASMIN hierarchy than in the case of the Hope–Goldthorpe prestige scores.

For example, class IIIa (clerical and routine non-manual workers) has nearly as much schooling as class II (lower professional and managerial workers), while class IV*d* (farmers without employees) has less schooling and less income than class VII*a* (semi-skilled and unskilled manual workers). In fact, if one considers the projection of each of the points in Figure 1 on the left-hand axes of the graphs, that is, the univariate distributions of classes by prestige, education, and income, it is clear in each case that the vertical dimension is finely graded and cannot adequately be specified by just two divisions among occupational strata. Still, it remains an empirical matter whether these more finely graded measures add anything to the explanation of mobility, beyond the vertical distinctions recognized in the CASMIN model. Before turning to that question, we first consider whether our disaggregation, however scaled, provides more information about the mobility process than the CASMIN scheme.

Heterogeneity of mobility

Two sets of occupational categories, one a more aggregated version of the other, can be said to provide the same information about mobility if the categories to be combined have homogeneous mobility rates, as indicated by independence in sub-tables of the more detailed mobility table. Two categories can be combined if the chi-square test is insignificant (Goodman 1981*a*). As shown in Tables 1 and 2, Erikson and Goldthorpe's combined categories do not have homogeneous mobility rates and should not be combined. The homogeneity tests proposed by Goodman (1981*a*) lead to a rejection of the null hypothesis of homogeneity for four of the five pairs of categories in all seven Western European nations (and all five pairs in three nations) and two of the three pairs of categories in the two Eastern European nations (see Table 1). For England and Wales, null hypotheses of homogeneous origins, homogeneous destinations, or both are rejected for four of the five pairs of occupational categories: the routine non-manual categories (III*a*,*b*), the proprietor categories (IV*a*,*b*), the farmer categories (IV*c*,*d*), and the upper manual categories (V, VI). For the professional and managerial categories (I, II), the individual tests fail to reject the null hypothesis of homogeneity by small amounts, but the combined test ($\Sigma L^2 = 30.69$; $df = 19$; $p < .05$) indicates rejection of homogeneity. For France, Germany, and Hungary, at least one test rejects homogeneity for each pair of occupational categories (five pairs for France and Germany; three pairs for Hungary because employers are not singled out). For Ireland, Northern Ireland, Scotland, and Poland, at least one test indicates rejection of homogeneity for four of five categories (two of three in the case of Poland). As in England and Wales, the exception is the distinction between upper and lower professionals and managers (I and II). For the three 'Celtic fringe' nations and Poland, the combined test does not indicate rejection, so the upper–lower professional and manager distinction is not significant for these four nations. In Sweden it is the distinction between farmers with and without employees (IV*c* and IV*d*) that is not significant in each test and in the combined test.

A global test of all five aggregations (three for Hungary and Poland) is the difference between the L^2 for independence in the full table and the L^2 for independence in the 7×7 table for the CASMIN categories (Goodman 1981*a*). Table 2 reports these difference tests.[5] For the Western European nations the difference of L^2s is distributed as chi-square with $(121 - 36 =)$ 85 degrees of freedom; for the Eastern European nations there are $(81 - 36 =)$ 45 degrees of freedom. In each of the nine nations, the difference of L^2s is highly significant. The minimum difference is for Sweden ($\Delta L^2 = 201.68$; $df = 85$; $p < .05$). Global tests like these are not very precise; using all 85 degrees of freedom at once is pretty crude. When a significant difference is observed, these tests do not reveal the source of heterogeneity. However, when coupled with the detailed tests in Table 1, the global tests in Table 2 provide an efficient summary of the magnitude of heterogeneity.

Chi-square tests applied to samples as large as the English, French, and Polish studies sometimes detect statistically significant differences that are substantively trivial. The ratio of the L^2s for independence in the CASMIN

TABLE 1 *Homogeneity of mobility for pairs of occupational categories: Men 25–64 years old, CASMIN data*

	Categories to be combined				
Component	I,II	III*a*,III*b*	IV*a*,IV*b*	IV*c*,IV*d*	V,VI
A. Western Europe					
England and Wales					
Destination	15.47[b]	20.98[a]	16.81	21.76[a]	20.73[a]
Origin	14.46[b]	34.65[a]	38.46[a]	15.87	69.83[a]
Diagonal	.76[b]	7.62[a]	.09	2.32	.36
France					
Destination	92.85[a]	35.63[a]	13.13	28.47[a]	32.66[a]
Origin	69.03[a]	7.40	85.74[a]	79.74[a]	67.48[a]
Diagonal	16.84[a]	<.01	.95	.50	5.03[a]
Germany, Federal Republic					
Destination	21.92[a]	17.02[a]	16.05	9.64	57.88[a]
Origin	11.51	12.32	17.96[a]	8.57	31.51[a]
Diagonal	10.71[a]	.88	19.21[a]	5.36[a]	8.21[a]
Ireland					
Destination	3.49	16.85	16.47	18.11[a]	28.67[a]
Origin	6.82	19.99[a]	18.08[a]	24.89[a]	20.41[a]
Diagonal	.35	1.40	2.48	1.43	5.12[a]
Northern Ireland					
Destination	4.77	17.30[a]	20.54[a]	5.23	2.53
Origin	13.68	6.48	9.60	33.50[a]	17.97[a]
Diagonal	.01	.43	2.12	.01	.88
Scotland					
Destination	8.30	27.28[a]	15.46	14.62	12.30
Origin	11.75	35.78[a]	27.87[a]	21.31[a]	21.31[a]
Diagonal	.12	.04	.44	9.64[a]	.03
Sweden					
Destination	14.45	13.52	7.93	4.55	18.09[a]
Origin	5.25	17.07[a]	20.38[a]	13.43	8.97
Diagonal	4.66[a]	1.95	.32	.18	.96
B. Eastern Europe					
Hungary					
Destination	18.23[a]	72.48[a]	—	—	19.31[a]
Origin	12.68	56.99[a]	—	—	13.92
Diagonal	.77	1.51	—	—	.41
Poland					
Destination	4.71	29.73[a]	—	—	23.62[a]
Origin	7.25	44.79[a]	—	—	43.86[a]
Diagonal	.26	.69	—	—	19.46[a]

Notes: [a]$p<.05$; [b]although these L^2 values are not significant at the .05 level, their sum is significant ($\Sigma L^2=30.69$; $df=19$; $p<.05$).

and full tables is an aid to interpreting L^2 difference tests. It measures the effectiveness of the aggregated CASMIN categories in representing the information available in the full set of categories. The last column of Table 2 reports this ratio for each nation. The CASMIN categories are most effective in Poland, where the two employer categories are not included in the full set of categories; the ratio of L^2s is 0.98. In Hungary where there are no employer categories either, the CASMIN categories are less effective; the ratio of L^2s is 0.88. France is the only Western European nation in which the effectiveness of the CASMIN set of

TABLE 2 *Likelihood ratio chi-square (L^2) for independence, quasi-independence, and quasi-symmetry models by occupational classification: Men 25–64 years old, CASMIN data*

	L^2			Percentage of baseline	
Country	Full	CASMIN	Difference	Full	CASMIN
A. Western Europe					
England and Wales (*N=8343*)					
Independence	2 396.65	1 984.72	411.93	100	83
Quasi-independence	1 118.70	445.62	673.08	47	19
Quasi-symmetry	107.10	47.85	59.25	4	2
France (*N=16 432*)					
Independence	8 766.92	7 778.12	988.80	100	89
Quasi-independence	3 941.36	1 551.87	2 389.49	45	18
Quasi-symmetry	207.40	68.78	138.62	2	<1
Germany (*N=3576*)					
Independence	1 586.46	1 203.05	383.41	100	76
Quasi-independence	686.10	268.62	417.48	43	17
Quasi-symmetry	121.43	35.72	85.71	8	2
Ireland (*N=1745*)					
Independence	1 378.21	1 098.33	279.88	100	80
Quasi-independence	547.31	192.97	354.34	40	14
Quasi-symmetry	68.77	14.23	54.54	5	1
Northern Ireland (*N=1807*)					
Independence	1 094.48	885.40	209.08	100	81
Quasi-independence	423.36	164.44	258.92	39	15
Quasi-symmetry	79.22	17.83	61.39	7	2
Scotland (*N=4066*)					
Independence	1 694.25	1 367.73	326.52	100	81
Quasi-independence	753.38	348.42	404.96	44	21
Quasi-symmetry	98.95	51.44	47.51	6	3
Sweden (*N=1880*)					
Independence	641.49	439.81	201.68	100	69
Quasi-independence	308.68	139.52	169.16	48	22
Quasi-symmetry	55.91	18.67	37.24	9	3
B. Eastern Europe					
Hungary (*N=10319*)					
Independence	2 995.80	2 646.98	348.82	100	88
Quasi-independence	1 950.77	1 556.02	394.75	65	52
Quasi-symmetry	76.75	50.67	26.08	3	2
Poland (*N=27 993*)					
Independence	9 408.57	9 192.56	216.01	100	98
Quasi-independence	1 945.61	897.53	1 048.08	21	10
Quasi-symmetry	125.29	90.84	34.45	1	<1

Note: Degrees of freedom are: (*a*) for the full set of categories in Western Europe, independence, 121; quasi-independence, 109; quasi-symmetry, 55; (*b*) in Eastern Europe, independence, 81; quasi-independence, 71; quasi-symmetry, 36; (*c*) for the CASMIN set of categories in all countries, independence, 36; quasi-independence, 29; quasi-symmetry, 21. Degrees of freedom for the differences between full and CASMIN can be obtained by subtracting the degrees of freedom for CASMIN from the degrees of freedom for full.

categories is as high as it is in Eastern Europe; the ratio of L^2s in France is 0.89. In the other Western European nations, the CASMIN set of categories masks from one-sixth to one-third of the association revealed by the full set of categories. In England and Wales, Ireland, Northern Ireland, and Scotland, the CASMIN set of categories are 17 to 20 per cent less effective than the full set of categories. In Germany the CASMIN set of categories is 24 per cent less effective than the full set, and in Sweden the CASMIN set misses nearly one-third of the association between origins and destinations.

Of course a sizeable portion of the overall association between origins and destinations in any mobility table is due to immobility. The data-sets assembled here are not exceptions to this generalization. Using the full set of categories (twelve in Western Europe and ten in Eastern Europe), the L^2 for quasi-independence is less than half of the L^2 for independence in eight of the nine nations. The extremes are Hungary, where immobility accounts for only 35 per cent of total association, and Poland, where immobility accounts for 79 per cent of total association.

The ineffectiveness of the CASMIN classification scheme would be much less consequential if the association that it fails to capture was primarily located along the diagonal of the smaller tables, i.e. if it was due to short-range mobility that appears as immobility in the CASMIN tables. The L^2 for quasi-independence using the CASMIN categories to that for quasi-independence in the larger table using the full set of categories measures the effectiveness of the aggregated categories in retaining the information available in the full set of categories.

If this appreciable immobility is the source of all of the lower effectiveness of the CASMIN categories, then the ratio of L^2s for quasi-independence for the CASMIN categories to that for the full set of categories will be 1.0. The further the ratio of L^2s from 1.0, the more important are other forms of mobility, particularly differential mobility from one category in a pair that is aggregated in the CASMIN scheme to another pair.[6] Table 2 shows the L^2s for quasi-independence for the full and CASMIN classifications, and their ratio to the baseline L^2 (that for independence in the full classification). In all of the nations except Hungary the L^2 for quasi-independence obtained when using the CASMIN categories is less than half that obtained when using the more detailed categories. In the Western European nations the ratio of L^2s ranges from 35 to 48 per cent.

In summary, the tests in Tables 1 and 2 show that the CASMIN classification scheme masks a substantial fraction of the total amount of differential mobility. More to the point, the association that is masked includes not only short-range mobility that is miscoded as immobility (a condition that might arguably be considered acceptable), but also an appreciable amount of differential mobility between and within aggregated categories.

SYMMETRICAL ASSOCIATION

The CASMIN model includes four asymmetries in its specification of the association between origins and destinations. First is the intrinsic asymmetry between farm origins and farm destinations. Erikson and Goldthorpe treat farm origins as being in the lowest stratum, but they treat farm destinations as being in the middle stratum. They base this specification on the larger average size of contemporary farms compared with farms that would have been typical in the father's generation. In addition, they specify that mobility from farm labour (class VII*b*) to lower blue-collar work (class VII*a*) will exceed the reverse flow from lower blue collar to farm labour by an amount greater than would be expected given their identical hierarchical level (the lowest), their sectoral difference, and the marginal effects.[7]

The asymmetry of farm origins and destinations strikes us as very strange indeed, for it implies that the prototypical form of class inheritance—the generational succession from father to son on a family farm—constitutes upward mobility. We understand that Erikson and Goldthorpe regard this as a compositional effect; they see the average size of destination farms being larger than the average size of origin farms and conclude that inheritance necessitates a certain size. Yet for the men actually involved in the inheritance there is no mobility. The mobility occurs for the ones who fail to inherit. Although neither they nor we have direct measures of the social or economic standing of farm fathers and sons, the CASMIN specification appears superficially weak because of the heterogeneous ages of men included in Erikson and Goldthorpe's analyses. That is, given the age-range from 20 to 64 that they use, there is a great degree of overlap between the temporal referents of father's and of son's occupations. To be sure, on average there is a historic difference of one generation between fathers and sons, but the range of ages of men in the CASMIN analysis (44 years) covers nearly twice the span of the typical human generation (about 25 years). In the CASMIN analysis, the class experiences of fathers and sons cover much the same historic period, and this again leads us to doubt the assumption by Erikson and Goldthorpe that the vertical position of farming has changed between generations.

The other proposed asymmetry—between labour on and off farms—seems to us more likely to be a function of marginal differences than of asymmetric association. In any event, symmetry is an empirical question, and all four asymmetries should be tested against quasi-symmetry, a model in which odds-ratios are completely symmetric about the main diagonal of the table, regardless of the ordering of the classes.

Table 2 reports the first phase of the test: the ratio of the L^2 for quasi-symmetry to that for independence in the full table for each nation. This ratio measures the asymmetric association (in the form of a residual from quasi-symmetry) as a fraction of total association. It is largest in Sweden, where 9 per cent of the total association is asymmetric; however, with 55 degrees of freedom, the L^2 of 55.91 is not significant, even though it is 9 per cent of the total association. In the other eight nations the L^2 for quasi-symmetry is statistically significant, but it is substantively trivial as 92–8 per cent of the total association is symmetric. Even if immobility is discounted by taking the L^2 for quasi-independence as the basis for comparison, 82–90 per cent of the off-diagonal association is symmetric.[8]

Perhaps more to the point is the comparison between the residual from quasi-symmetry for the CASMIN categories and independence for the full set of categories, since it is the CASMIN categories that are treated asymmetrically by Erikson and Goldthorpe. In each nation the asymmetry of association in the CASMIN categories accounts for 3 per cent or less of the total association contained in the full classification. In the most extreme cases, the test statistic for asymmetry in Scotland is 51.44, which is just over 3 per cent of the L^2 for independence in the full table; in France, the test statistic for asymmetry is 68.78, which is less than 1 per cent of the L^2 for independence in the full table. Especially in light of the advantages to be gained in the interpretation of structural mobility from the assumption of symmetry, these findings do not make a strong prima facie case for the asymmetric design of the CASMIN model.

The case against asymmetry is not completely cut and dried. The L^2 for quasi-symmetry is significant in eight of the nine nations. However, in each nation the value of Raftery's (1986*a*, 1986*b*) approximation to *bic* is negative, indicating that quasi-symmetry is preferred to the asymmetric saturated model by Bayesian criteria of inference. Of course, it might be possible to find an acceptable asymmetric alternative that is less complicated than the saturated model (the basis of comparison when using *bic*), but that contains one, two, or more asymmetries. We have not gone further in ransacking the data to locate possible asymmetries, but we have tested the four asymmetries proposed by Erikson and Goldthorpe. We begin with quasi-symmetry for the CASMIN categories, parameterized according to the model proposed by Sobel *et al.* (1985). To this parameterization we add four dummy variables: $\epsilon_{61}=1$ if origin is in class (IV*c*,IV*d*), destination is in class (I,II) and 0 otherwise; $\epsilon_{62}=1$ if origin is in class (IV*c*,IV*d*), destination is in classes III, (IV*a*,IV*b*), V, or VI, and 0 otherwise; $\epsilon_{65}=1$ if origin is in class (IV*c*,IV*d*), destination is in class VII*a*, and 0 otherwise; $\epsilon_{75}=1$ if origin is in class VII*b*, destination is in class VII*a*, and 0 otherwise. Table 3 reports the goodness-of-fit statistics

TABLE 3 *Goodness-of-fit statistics for quasi-symmetry and asymmetric models of association for the CASMIN categories: Men 25-64 years old, CASMIN data*

Model	L^2	X^2	*df*	*bic*
A. Western Europe				
England and Wales				
Quasi-symmetry	47.85	64.21	15	−88
Asymmetry	20.44	21.88	11	−79
Difference	27.41	—	4	−9
France				
Quasi-symmetry	68.78	79.19	15	−77
Asymmetry	23.70	23.80	11	−83
Difference	45.08	—	4	6
Germany				
Quasi-symmetry	35.72	34.21	15	−87
Asymmetry	30.89	28.92	11	−59
Difference	4.83[a]	—	4	−28
Ireland				
Quasi-symmetry	14.23[a]	14.62[a]	15	−98
Asymmetry	5.98[a]	5.98[a]	11	−76
Difference	8.25[a]	—	4	−22
Northern Ireland				
Quasi-symmetry	17.83[a]	18.55[a]	15	−95
Asymmetry	8.61[a]	7.71[a]	11	−76
Difference	9.22[a]	—	4	−21
Scotland				
Quasi-symmetry	50.44	70.95	15	−74
Asymmetry	16.21[a]	16.69[a]	11	−75
Difference	34.23	—	4	1
B. Eastern Europe				
Hungary				
Quasi-symmetry	50.67	73.80	15	−88
Asymmetry	26.91	27.91	11	−75
Difference	23.76	—	4	−13
Poland				
Quasi-symmetry	90.84	96.85	15	−63
Asymmetry	58.79	61.33	11	−51
Difference	32.05	—	4	−12

Note: [a]Not significant at the .05 level.

obtained for quasi-symmetry and for quasi-symmetry augmented by these four asymmetries. In five of the eight nations the augmented model significantly improves fit over that obtained for quasi-symmetry. In three of these, however, $bic > 0$ indicates that quasi-symmetry should be preferred. Only in France and Scotland is the asymmetric model preferred.

The significant asymmetries differ in France and Scotland. For France, the parameter estimates (with standard errors in parentheses) are: $\epsilon_{61} = -1.078$ (.268), $\epsilon_{62} = -.072$ (.106), $\epsilon_{65} = .620$ (.210), and $\epsilon_{75} = .822$ (.217); for Scotland they are: $\epsilon_{61} = -2.574$ (.514), $\epsilon_{62} = -.871$ (.279), $\epsilon_{65} = .737$ (.784), and $\epsilon_{75} = .262$ (4.07). In France, mobility from farmer origins to professional destinations is less than would be expected based on symmetrical association, and mobility from farmer or farm-worker origins to a lower manual destination is greater than would be expected on the basis of symmetry. In Scotland ϵ_{61} is also significant, but ϵ_{65} and ϵ_{75} are not significant while ϵ_{62} is.

These asymmetries in France and Scotland may be worthy of further discussion. But asymmetries are clearly less general than is assumed by Erikson and Goldthorpe. Three of the four asymmetries they propose are significant in only one nation each. Together the three significant asymmetries in France contribute less than 1 per cent of the total association between origins and destinations in that nation. In Scotland, the significant asymmetries account for 2 per cent of the total association (34.23/1,694.25 = .02); other, unspecified asymmetries account for the remaining 1 per cent (16.21/1,694.25 = .01).

We have carried out one further test of the CASMIN specification that the location of farming in the vertical hierarchy has changed between generations. In essence, we compare the position of each of the 12 occupational classes among fathers with its position among sons. We specify the ranks of classes by fitting association models (Goodman, 1984) to the seven Western European mobility classifications in which all 12 classes can be distinguished. In the first of these models (line 7 of Table 4), we specify a completely free scaling of the classes, subject to the constraints that scale values must be the same in each nation and that scale values must be the same for father's class as for son's class.[9] This model yields an L^2 of 3,499 on 824 degrees of freedom, for which $bic = -5{,}187$. In the second model (line 6 of Table 4), we permit the row and column scale values to differ between father's class and son's class, but not among nations. This model yields an L^2 of 3,183 on 814 degrees of freedom, for which $bic = -5{,}397$.[10] Both models fit satisfactorily (using the $bic < 0$ criterion), but the latter model—with different class positions for fathers

TABLE 4 *Goodness-of-fit statistics for select models for 12 × 12 × 7 tables: Men 25–64 years old, Western European nations in the CASMIN data (N = 37 849)*

# Model	L^2	χ^2	*df*	*bic*	Δ
0 X*C + Y*C	17 557	23 679	847	8629	24.3
1 X*C + Y*C + X*Y	1 457	1 491	726	−6196	5.6
2 X*C + Y*C + SYM	1 758	2 023	781	−6475	6.3
3 X*C + Y*C + SYM*C	739	954	385	−3319	2.9
4 X*C + Y*C + ALL	3 275	3 388	811	−5 274	10.0
5 X*C + Y*C + HI1 + HI2 + IN1 + IN2 + IN3 + SE1 + AF1 + AF2	3 416	3 602	839	−5428	10.3
6 X*C + Y*C + DIAG + DIAG.X + {Homogeneous RC}	3 183	5 750	814	−5397	8.7
7 X*C + Y*C + DIAG + DIAG.X + {Homogeneous equal RC}	3 499	4 278	824	−5187	9.6
8 {Model 7} + DIAG.C	3 462	4 214	818	−5161	9.4
9 {Model 7} + PHI.C	3 453	4 609	818	−5169	9.5
10 {Model 7} + DIAG.C + PHI.C	3 408	4 617	812	−5152	9.2
11 X*C + Y*C + DIAG + DIAG.X	8 083	10 638	835	−719	14.8
12 {Model 11} + SES + PRES	5 021	7 735	833	−3760	10.0
13 {Model 11} + SES + PRES + AG + SE2 + AG.SE2	2 455	2 658	830	−6294	8.0
14 {Model 11} + HI1 + HI2 + AG + SE2 + AG.SE2	2 790	2 887	830	−5959	9.1
15 {Model 13} + DIAG.C + SES.C + PRES.C + AG.C + SE2.C + AG.SE2.C + DIAG.X.C	2 212	2 447	794	−6157	7.6
16 {Model 13} + DIAG.C + SES.C + PRES.C + AG.C	2 252	2 501	806	−6244	7.7

Notes: All test statistics significant at $p < .05$; The symbols are: X = origin (coded as a twelve-category factor); Y = destination (coded as a twelve-category factor); C = nation (coded as a seven-category factor); SYM = an exhaustive set of interaction effects that are symmetric about the main diagonal; ALL = a set of interaction effects coded to represent all interactions in Erikson and Goldthorpe's 7 × 7 collapse of the 12 × 12 classification; HI1 = 1 for cells designated as moves of one hierarchical level in the CASMIN model and 0 otherwise; HI2 = 1 for cells designated as moves of two hierarchical levels in the CASMIN model and 0 otherwise; IN1, IN2, IN3 = diagonal blocks as coded in the CASMIN model; SE1 = the sector variable as coded in the CASMIN model; AF1, AF2 = affinities as coded in the CASMIN model; DIAG = 1 for cells on the main diagonal of the 12 × 12 classification and 0 otherwise; PHI = linear-by-linear interaction of freely scaled scores for origin and destination categories; PRES = a linear-by-linear interaction of prestige (the product of scores on the Hope-Goldthorpe scale for the origin and destination) for each cell; SES = a linear-by-linear interaction of occupational socio-economic status (based upon Ganzeboom *et al.*); AG = a dummy variable for agricultural sector employment coded as 1 for any combination of classes in the agricultural sector (IV*c*, IV*d*, VII*b*) and 0 otherwise; SE2 = a dummy variable for self-employment coded as 1 for any combination of classes involving substantial self-employment (I, IV*a*, IV*b*, IV*c*, IV*d*) and 0 otherwise; RC = Goodman's log-multiplicative row × column model.

and sons—fits somewhat better. Aside from effects of class inheritance, these models attempt to fit the CASMIN tables using a single dimension of class position, pertaining to origins, destinations, or both.

How different are the estimated scale values for father's class and son's class (origin and destination effects), and how do they differ? Figure 2 shows the scale values for father's class and for son's class under the second model, plotted in relation to the common values estimated under the first model. That is, the line in Figure 2 gives the scale values under the constraint of equality between class scales of fathers and sons, and the points show the actual scale values of father's class or son's class. First, there is little change across generations in the estimated positions of classes. The limited scatter of the points about the line shows that origin and destination scale values are each similar to the pooled scale values. Second, the largest inter-generational change in class position is a reversal between classes II and I, which are aggregated in the CASMIN model. In addition, there are smaller inversions between classes IV*a* and III*a*, and between classes V and

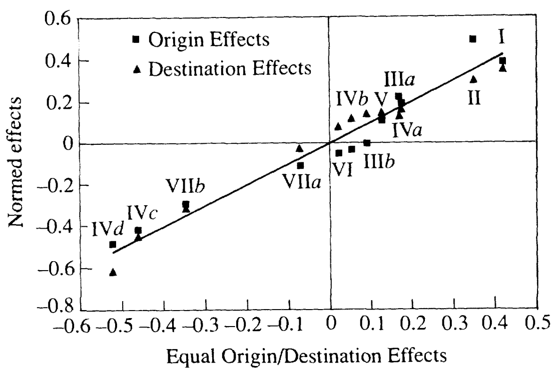

FIGURE 2 *Origin and destination effects by equal origin-destination effects: 12 × 12 mobility tables for 7 nations.*

III*b*, where the first member of each pair ranks higher as an origin category than as a destination category. But there are no inversions of more than one position in the rank order of scale values. Third, there is no evidence of any change in the standing of farm occupations relative to other occupational classes. Among sons as among fathers, farm occupations are at one extreme of the occupational scale, far distant from all other occupations. That is, the patterns of movement among classes offer no support for the hypothesis that the relative standing of farm occupations differs between the generations of fathers and sons in the CASMIN data.

STRUCTURAL MOBILITY

Symmetric association is attractive because of its parsimony and interpretability. In the absence of marginal differences between the distributions of origins and destinations, mobility flows from one class to another would be compensated by equal flows in the opposite direction. With asymmetry, marginal differences are intrinsic to the structure of association; they ascribe the dissimilarity between the origin and destination distributions to a combination of factors, some of them independent of origins, some of them origin-specific. As long as association is asymmetric, it makes no sense to talk about structural mobility from origin to destination 'in the absence of marginal differences' because the model implies that there will be marginal differences as long as there are asymmetries.

Intrinsic asymmetry is incompatible with the analysis of structural mobility. On the other hand, if quasi-symmetry holds, then all discrepancies between origins and destinations can be attributed to factors that are independent of origins (Sobel *et al.*, 1985). Given the substantial symmetry of the CASMIN mobility tables, we present estimates of origin-specific structural mobility for each nation. While prior, non-parametric approaches to structural mobility typically specified one structural mobility measure for a given table, the approach proposed by Sobel *et al.* (1985) differentiates growth in some categories and decline in others, presenting positive and negative structural mobility multipliers for each origin.

Figure 3 graphs the structural mobility multipliers (in log form) for each combination of occupational origin and nation. Log form is used to aid interpretation: $\log(\alpha_j) > 0$ indicates growth and $\log(\alpha_j) < 0$ indicates decline of occupation j. All sub-parts of the figure use the same scale, although the entire range of values for origins in a household headed by a farmer without employees (which is the origin with the greatest declines) is graphed on the negative side of the scale while the others are equally divided into positive and negative values.[11]

Structural mobility varies more from origin to origin than it does from nation to nation. The largest component of structural mobility in every nation except Scotland is the decline of small farming (IV*d*). The nation-to-nation variance in structural mobility is greatest in this category, ranging from a negative shift of −5.51 in Hungary to the neighbourhood of −3.25 in Sweden, Northern Ireland, and Germany, to the neighbourhood of −2.50 in Poland, Ireland, and France, to a low around −1.50 in England and Wales and in Scotland.[12] There are also large negative shifts in the other two farm origin categories, except among German farm-workers (an extremely small origin category).

The largest and most consistent growth categories are the professions and management (I and II). Hungary shows higher than average growth in these categories;[13] France shows

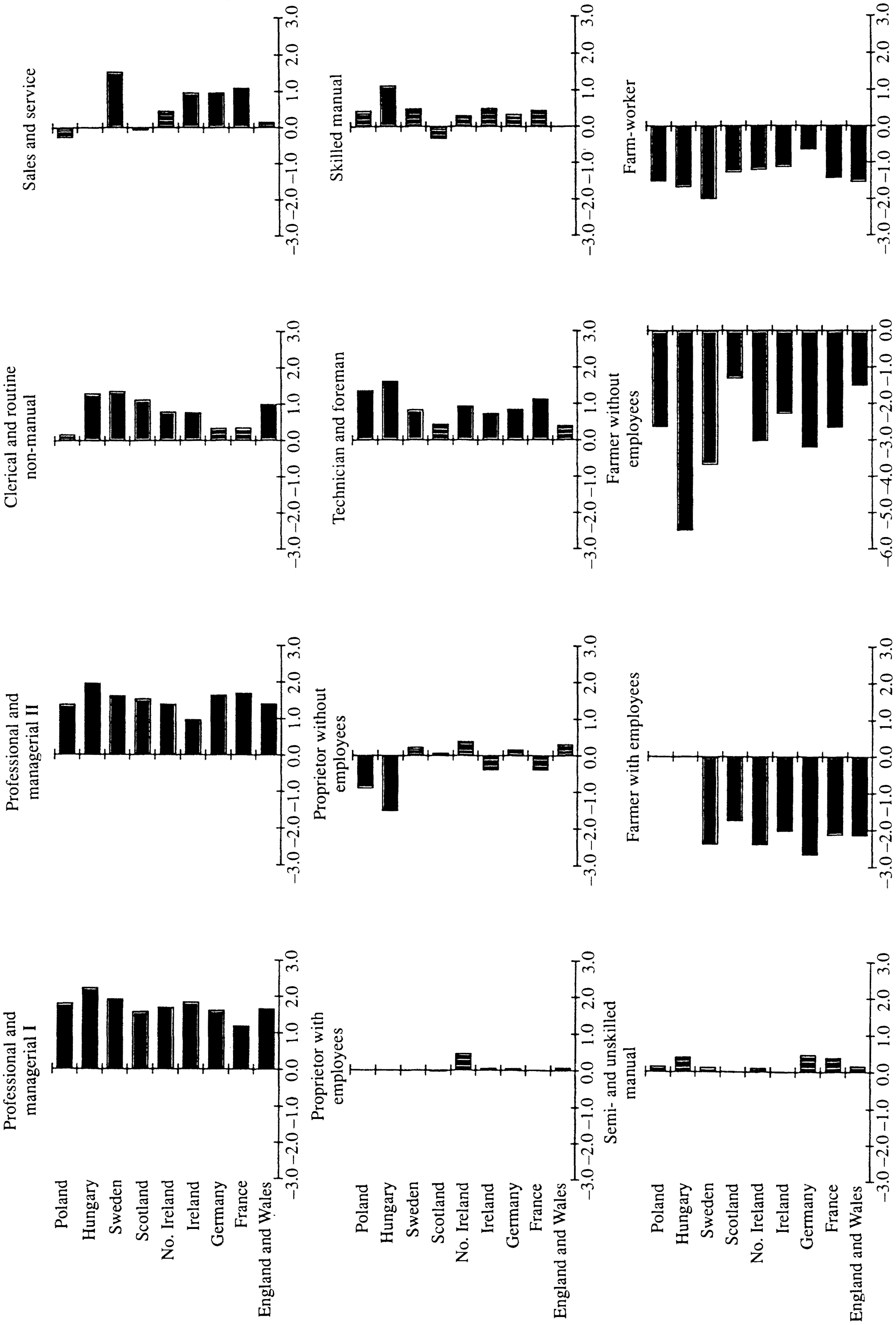

FIGURE 3 *Structural mobility multipliers (in log form) for European countries: men 25–64 years old, CASMIN data.*

lower than average growth in higher grade professionals and managers while Ireland shows lower than average growth in the lower grade professionals. The clerical and routine non-manual (III*a*), the sales and service (III*b*), and the technicians and foremen (V) categories show moderate growth in most nations. Exceptions include insubstantial positive shift for clerical and other routine non-manual workers in France, Germany, and Poland, for sales and service workers in the UK (England and Wales, Scotland, and Northern Ireland), Hungary, and Poland, and for technicians and foremen in England and Wales and in Scotland. For most nations, the structural mobility coefficients show near stasis for proprietors (IV*a* and IV*b*) and manual workers (VI and VII*a*). Exceptions are the decline of self-employment in the two Eastern European nations, and a positive shift of 1.13 for skilled manual work in Hungary.

In summary, factors that are independent of social origins redistributed workers away from agriculture and into post-industrial and technical employment in all nations. The nation with the greatest overall redistribution due to structural mobility is Hungary,[14] followed by Sweden. The least overall redistribution due to structural mobility is found in Scotland and in England and Wales.

MODELLING THE HIERARCHICAL DIMENSION

Erikson and Goldthorpe assign occupational categories to hierarchical levels and specify that mobility between any pair of categories is an inverse function of the number of levels separating them. Given this formulation, they find that hierarchical effects are much weaker than commonly supposed (e.g. Hope, 1981, 1982; Hauser, 1984*a*, 1984*b*; Hout, 1984*a*; Hout and Jackson, 1986). They have understated the case for hierarchical effects in three ways: (1) their aggregated categories erase significant hierarchical distinctions between pairs of categories; (2) their three hierarchical levels are insufficient to capture the relevant vertical distinctions among occupational categories, even within the context of their aggregated occupational classification; and (3) the 'social distance' and 'crossing parameters' specifications that they combine are inferior to a linear-by-linear association model for theoretical and empirical reasons.

The aggregated categories mask a significant portion of the overall association between origins and destinations, according to Goodman's (1981*a*) homogeneity criteria, as shown in Table 1 above. If the masked association is due to hierarchical effects, as is likely, then Erikson and Goldthorpe's estimates of modest hierarchical effects understate the true effects.

Simplistic hierarchies like Erikson and Goldthorpe's betray the continuous differentiation of occupations along the entire range of social standing from low-skill, low-paying occupations to high-skill, high-paying ones (Duncan, 1961; Hodge, 1963, 1981; Featherman and Hauser, 1978; Stevens and Cho, 1985). The objective continuum is echoed in the fine grain of subjective assessments typical of prestige studies; people are capable of distinguishing among closely related occupations (e.g. Hodge *et al.*, 1966; Goldthorpe and Hope, 1974; Treiman, 1977; Jencks *et al.*, 1988). Fine-grained distinctions have substantively important consequences in England and Wales, France, and Sweden (Hauser, 1984*a*, 1984*b*), in Scotland (Hope, 1986), in Ireland and Northern Ireland (Hout, 1989), and in the United States (Duncan and Hodge, 1963; Blau and Duncan, 1967; Hout, 1984*a*, 1988). Specifying too few hierarchical levels is similar to using too few occupational categories in that it leads to an underestimate of the magnitude of the true hierarchical effect(s).

The CASMIN model uses two design matrices labelled HI1 and HI2 to operationalize hierarchical effects. The design matrices combine the formal properties of social distance models and crossing models (Goodman, 1972; Hout, 1983). Erikson and Goldthorpe's (1987*a*) discussion makes it plain that they intend the social distance formulation.

The choice between a social distance and a linear-by-linear specification is a technical matter that at first seems somewhat arbitrary. Upon examination it becomes clear that the implications of the two models are dramatically different. The linear-by-linear formulation accords more closely with the prevalent conceptualization of stratification processes, and we advocate its use in future work with the CASMIN data and other data.

The social distance model specifies that the amount of mobility between origin i and destination j is an inverse function of the number of levels that separate i and j:

$$F_{ij} = \mu\alpha_j\beta_i\beta_j\gamma^{|X_i - X_j|} \quad (1)$$

where F_{ij} is the expected frequency in cell (i, j), $\mu > 0$, $\alpha_j > 0$, $\beta_j > 0$, $\gamma > 0$,[15] μ is the 'grand mean' term, the α_j are 'structural mobility multipliers' (Sobel *et al.*, 1985), β_i are 'composition multipliers,' $\beta_i = \beta_j$ if $i = j$, $\Pi_j\alpha_j = \Pi_j\beta_j = 0$, γ is the association parameter, and $X_i - X_j$ is the distance between row i and column j measured in number of levels or on some other scoring metric (e.g. number of prestige-scale points). In most empirical cases, $\gamma < 1$, indicating that mobility between classes i and j decreases as distance $(X_i - X_j)$ increases. If we let $D_{ij} = |X_i - X_j|$, then the logit form of the distance model is:

$$\begin{aligned} Y_{ijj'} &= \log(F_{ij}/F_{ij'}) \\ &= \lambda_0 + \lambda_1(D_{ij} - D_{ij'}) \end{aligned} \quad (2)$$

where $\lambda_0 = \log(\alpha_j\beta_j/\alpha_{j'}\beta_{j'})$ and $\lambda_1 = \log\gamma$. For any i, $(D_{ij} - \mathrm{D}_{ij'})$ can have (at most) two values, $-k_{ijj'}$ and $k_{ijj'}$. If we set $k_{ijj'} \geqslant 0$ and order the i so that $X_1 \geqslant X_2 \geqslant \ldots \geqslant X_R$ (and order the j the same way), then:

$$\begin{aligned} (D_{ij} - D_{ij'}) &= -k_{ijj'} \; \textit{for } i \leqslant j, \\ &= k_{ijj'} \text{ otherwise.} \end{aligned} \quad (3)$$

Together (2) and (3) imply that the logistic regression of Y_{ijj}' on i is a simple step function with at most two values (if $\lambda_1 = 0$ or $k_{ijj'} = 0$, then $\mathrm{Y}_{ijj'} = \lambda_0$, a constant). The left half of Figure 4 depicts this one-step relationship between origins and destinations according to the social-distance model.[16]

Thus the social-distance model provides an undifferentiated picture of the way that background affects achievement. It says that origin effects are limited to a single increment to the odds on a good job for the offspring. That single effect takes the form of odds on immobility and downward mobility that are higher than the odds on upward mobility. Within those two broad groupings of origins the odds on a better job are identical.[17] This lack of differentiation across most of the range of origins is not an acceptable specification of key theories of intergenerational stratification (Heath, 1981: 11–47). Implicit or explicit in formulations from Sorokin to Blau–Duncan to Parkin is the notion that, however occupational distinctions are conceived, increments to the social standing of an individual's occupation will result in advantages in the second generation. The social-distance model implies substantially less differentiation of life-chances than these and other key theorists and researchers suppose.

In contrast, the linear-by-linear model specifies the incremental advantages that most theorists and researchers associate with stratification. The linear-by-linear model specifies that the amount of mobility between origin i and destination j is a positive function of the scores for categories i and j (whether those scores are simply integers for the levels or some other metric such as prestige):

$$F_{ij} = \mu\alpha_j\beta_i\beta_j\theta^{X_iX_j} \quad (4)$$

where F_{ij}, μ, α_j, and β_j, are defined as in (1), θ is the association parameter, and X_i is the score for occupational category i (Duncan, 1979; Goodman, 1979*a*). In most empirical cases, $\theta > 1$, indicating that mobility between classes i and j increases as the product X_iX_j increases.[18] In this form it is not obvious why the linear-by-linear model accords better with stratification theory than the social-distance model does. The logit form of the linear-by-linear model makes its greater suitability clear:

$$Y_{ijj'} = \lambda_0 + \lambda_1 X_i \quad (5)$$

where $\lambda_0 = \log(\alpha_j\beta_j/\alpha_{j'}\beta_{j'})$ and $\lambda_1 = \log\theta\,(X_j - X_{j'})$. As (5) makes clear, the linear-by-linear model specifies that the log-odds on a better job are a linear function of the origin score. The right half of Figure 4 depicts this linear relationship between origins and destinations.

The comparison between the models in (1) and (4) can also be made in a dramatic way using

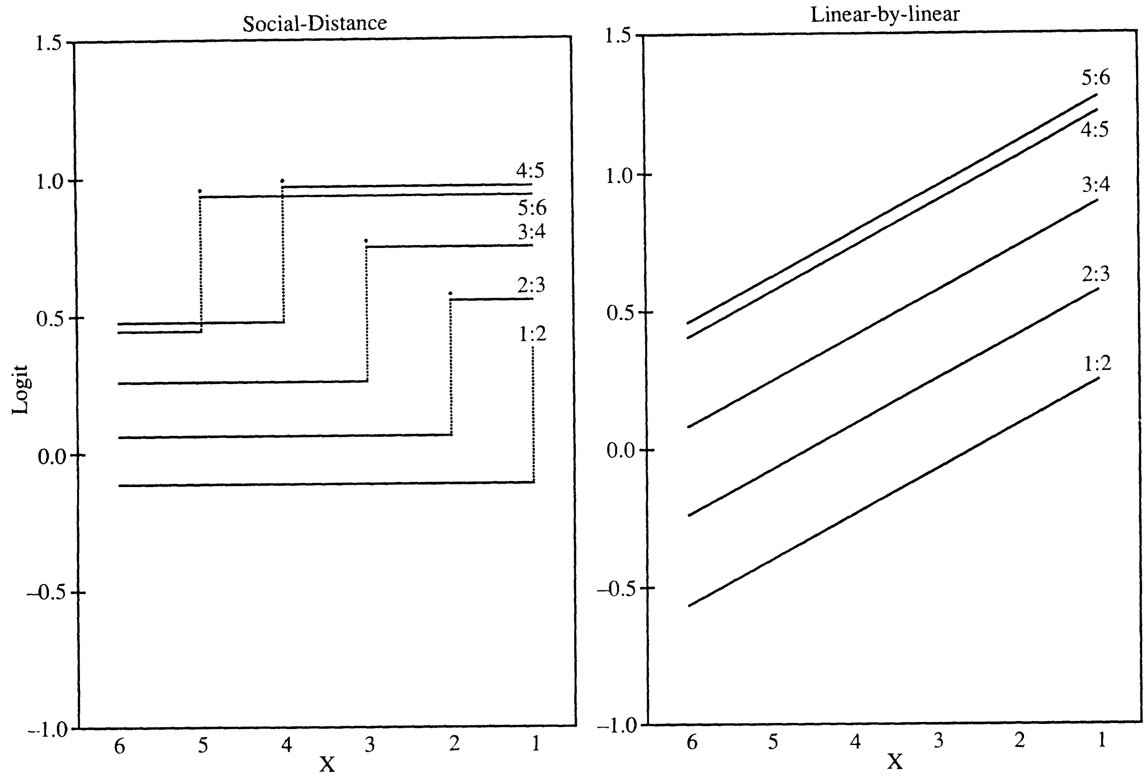

FIGURE 4 *Relationship between logit and origin for social-distance and linear-by-linear dependence models.*

the corresponding odds-ratios (see Goodman, 1979*b*). If we order the i so that $X_1 \geqslant X_2 \geqslant \ldots \geqslant X^R$ (and order the j the same way), as we did for (3), and consider only the 'minimal set' of odds ratios for adjacent categories then the social distance model (1) implies:

$$\begin{aligned} \theta_{ij} &= \gamma^{2(X_i - X_{i+1})} \text{ if } i=j; \\ &= 0 \qquad\qquad\quad \text{otherwise.} \end{aligned} \tag{6}$$

i.e. the social-distance model implies quasi-independence away from the diagonal. On the other hand, the linear-by-linear model (4) implies non-zero association throughout the minimal set of odds ratios (except for adjacent categories that have identical scores):

$$\theta_{ij} = \theta^{(X_i - X_{i+1})(X_j - X_{j+1})}. \tag{7}$$

(1) seems to accord well with notions about mobility in that most analysts would agree, on the face of it, that the volume of mobility decreases as the social distance between categories increases. But stratification theory is more fundamentally concerned with how destinations depend on origins. (2) and (5)—the logistic regression forms of the social-distance and linear-by-linear models—are the appropriate expressions of dependence relationships (Goodman, 1981*b*). From these two equations it appears that the linear-by-linear model captures the relationship between origins and destinations better than the social-distance model does. (2) and (3) show that the social-distance model (1) implies, contrary to most understandings, that advantages of origin are not uniform; they only accrue as safeguards against downward mobility.[19] (6) provides another view of the lack of effect implied by the social-distance model.

The attractive conceptual aspects of the social-distance formulation can be reconciled with the better modelling of dependence relations in the linear-by-linear formulation by squaring social distance instead of taking its absolute value as in (1). The squared social-distance model is equivalent to the linear-by-linear model, i.e. they have identical expected frequencies (Hope, 1981; Goodman, 1991).

This formal analysis helps us understand what is at stake in choosing between the linear-by-linear and social-distance formulations. Ultimately, it is necessary to turn to the data in order to adjudicate differences, recognizing that the data may at times be equivocal as to the proper formulation (e.g. Hauser 1984*a*, 1984*b*). In particular, the use of design matrices such as HI1 and HI2 minimally restricts the statistical search for fit, making it both difficult to reject and difficult to interpret (Hout, 1989: 148–52).

RE-ANALYSIS OF THE CASMIN DATA

Baseline models

We have faulted the CASMIN model because its categories are too highly aggregated, it includes unnecessary asymmetric effects, it collapses the hierarchical distinctions among classes into only three levels, and it is based on the social-distance model. We redress these shortcomings in the CASMIN model by using the full set of twelve occupational categories, a symmetric, linear-by-linear functional form, and continuous variables to distinguish social standing among classes.

To facilitate the presentation of models, we switch here to a GLIM-style notation, i.e. we drop subscripts and coefficients. The unit of analysis is the cell of the origin × destination × nation table; the link function is log, and the error structure is Poisson (McCullagh and Nelder, 1983). The baseline model is:

$$\log F = X^*C + Y^*C \tag{8}$$

where F = the expected frequency in cell $\{i,j,k\}$, X = origin (coded as a twelve-category factor), Y = destination (coded as a twelve-category factor), C = nation (coded as a seven-category factor), and the '*' operator fits the full set of marginals implied by the variables it joins. The model in (8) is a 'conditional independence' model (Goodman, 1970) in that it fits the two-way marginals of origin-by-nation and destination-by-nation, but it specifies that, within nation, origins and destinations are independent. As shown in Table 4 (line 0), the CASMIN data lead us to reject this simplistic model ($L^2 = 17{,}557$; $df = 847$; $bic = 8{,}628$). The data are fitted well by another baseline model (line 1 of Table 4)—referred to as the 'common social fluidity model' by Erikson and colleagues (1982)—which specifies cross-nationally invariant, but otherwise unconstrained association between origins and destinations ($L^2 = 1{,}457$; $df = 726$; $bic = -6{,}196$). One might take the fit of this model as a standard relative to other, more restrictive models specifying cross-nationally invariant association between origins and destinations. However, in light of our findings about symmetry in the mobility process, we also introduce a restricted model of cross-nationally invariant social fluidity, namely, constant quasi-symmetric mobility (CQS). As shown in line 2 of Table 4, CQS fits almost as well as the model of constant social fluidity ($L^2 = 1{,}758$; $df = 781$; $bic = -6{,}475$), and, using *bic* as the criterion, its fit is better than that of any other model we have considered, including the CASMIN model (line 5 of Table 4). Unfortunately, this model is substantively unsatisfactory, because it is based upon a completely arbitrary specification of origin-by-destination interactions, subject only to the restrictions of quasi-symmetry and cross-national invariance. Thus, we have sought to find more interpretable models whose fit to the CASMIN tables resembles that of CQS.

In line 3, we report the fit of the model of variable quasi-symmetry, which tells us the best fit that could be obtained under any model that does not violate the constraints of quasi-symmetry. Model 3 is simply the combination of the several nation-specific models of quasi-symmetry, whose fit is reported in the first column of Table 2 and whose parameters were used in our analyses of structural mobility. While the chi-square test statistics are very small under this model, so are its degrees of freedom, and, based upon the *bic* statistic, it is less desirable than other models we consider for the combined set of West European mobility classifications.

Models 4 and 5 of Table 4 provide benchmark measures of the fit of the CASMIN model, which are useful in later comparisons. Model 4 fits an interaction effect for every distinct combination of rows and columns in the 7 × 7 CASMIN tables. That is, it projects a saturated model of origin-by-destination interaction in the 7 × 7 CASMIN table on to the disaggregated, 12 × 12 table. The test statistics for this model tell us the amount of association in the 12 × 12 tables that could not possibly be explained by a cross-nationally invariant model of the 7 × 7 CASMIN tables, including cross-nationally variable association and association within cells that are collapsed in the 7 × 7 tables. This model fits well, in the sense that, in the absence of other alternatives, we would not reject it using *bic* ($L^2 = 3{,}275$; $df = 811$; $bic = -5{,}274$), but, like the arbitrary parameterizations of constant social fluidity or of quasi-symmetry in the full table, it lacks substantive appeal.

Fitting the CASMIN model

In line 5 we report the fit of the CASMIN model within the seven Western European nations for which we can analyse the 12 × 12 tables. Its fit statistics are almost as good as those of model 4, which is to say that the CASMIN model accounts very well for cross-nationally invariant association in the 7 × 7 tables. Since the CASMIN model uses fewer parameters than model 4, its *bic* statistic is preferable (more negative). While we will use the fit of the CASMIN model as one of our points of comparison, we note that, relative to other models, the CASMIN model is privileged by having been tailored to the data at hand. It will probably be more useful to validate the CASMIN model against fresh data than to compare any measure of its fit with that of other models of the same data.

For example, the international class-structure project (Wright, 1985), replicated in Britain in 1984, provides an opportunity for limited cross-validation (Marshall *et al.*, 1988). Three caveats limit the value of this replication: (1) the comparison between 1972 and 1984 allows for substantial true change in the population, (2) the relatively small number of cases (N = 659 men 25–64 years old) in the data used by Marshall and colleagues makes the statistical power of the replication relatively low, and (3) the absence of men with destinations in class III*b* and the lack of information on employees among fathers who were farmers makes it impossible to make the distinctions between III*a*/III*b* and IV*c*/IV*d*.

We begin the cross-validation by fitting the CASMIN model to the 10 × 10 British mobility table for 1972 (combining III*a*/III*b* and IV*c*/IV*d*). The model does not fit by conventional criteria, but *bic* shows it to be preferable to the saturated model on Bayesian grounds ($L^2 = 417.55$; $df = 73$; $bic = -242$). The parameter estimates (see Table 5) repeat the pattern familiar to those who have worked with the CASMIN model. The replication produces mixed evidence on the validity of the CASMIN model. The model fits the 1984 data quite well ($L^2 = 80.83$; $df = 73$; $bic = -393$).[20] More importantly, though, several of the parameter estimates are wildly different (see Table 5): the effect of IN1 is too small (and statistically insignificant) and those of HI2 and SE are too large (although not significantly different from the 1972 effects in a two-tailed test).

The key differences are in the affinity terms. We regard it as noteworthy that these are substantively the weakest terms in the model; they are less grounded in theory and prior research than inheritance, hierarchy, or sector. The 'negative affinity' term (AF1) actually has a large *positive* coefficient as well as a very large standard error because it is based on the contrast between one case and the other 658 cases. So we dropped the affinity terms (AF1 and AF2) from the model for the 1984 data. Simplifying the model in this way actually improves *bic* slightly while increasing L^2 by a barely significant amount ($L^2 = 90.20$; $df = 75$; $bic = -397$). The parameter estimates from this simpler model of the 1984 data reproduce the findings in the 1972 data better than do those in the full CASMIN model (see Table 5). Dropping the affinity terms produces a smaller HI2 coefficient with a much smaller standard error. However, the estimated effects of IN1 and SE remain problematic.

TABLE 5 *Parameter estimates by year for the CASMIN model for 10×10×2 tables: Men 25–64 years old, Great Britain, 1972 and 1984*

	1972		1984		1984	
Term	b	s.e.	b	s.e.	b	s.e.
Inheritance						
IN1	.360	.041	.099[a]	.160	−.029[a]	.159
IN2	.570	.070	.663	.231	.773	.230
IN3	1.218	.238	1.287[a]	1.104	1.175[a]	1.099
Hierarchy						
HI1	−.393	.034	−.455	.117	−.347	.109
HI2	−.502	.062	−1.211	.749	−.546	.173
Sector						
SE	−1.246	.088	−3.014	1.211	−3.082	1.206
Affinity						
AF1	−.712	.213	.888[a]	.762	—	—
AF2	.428	.033	.335	.119	—	—

Note: [a]Not significant at the .05 level.

In sum, the data from Marshall and colleagues lack sufficient statistical power to be definitive in cross-validating the CASMIN model. All in all, while there is not a failure of replication, neither are we impressed by the similarity in findings. The differences between the original British data and the 1984 replicate may have arisen by chance or from changes in the process of stratification in Britain, but our admittedly sceptical reading is that they reflect problems with the CASMIN model, especially the affinity parameters.

Alternative models

Had we no other alternatives, the fit of models 4 and 5 would be acceptable, but they do not fit the full 12×12 classification as well as either model 1 or model 2. Constant social fluidity and constant, quasi-symmetric social fluidity both yield superior fit, whether we look at the chi-square test statistics or at *bic*. Despite the several theoretical and conceptual flaws that we have identified in the CASMIN model, its only claim to superiority over models 1 or 2 is its theoretical content, and even that claim is weakened by inclusion of the affinity parameters.

As one relevant point of comparison to the CASMIN model, we consider two of Goodman's log-multiplicative association models (Goodman, 1984), whose fit is shown in lines 6 and 7 of Table 4. In model 6, homogeneous row and column effects, we choose scale values for origins (row effects) and destinations (column effects) so as to maximize the explained association in the tables, subject to the constraints that the row effects are the same in each nation and the column effects are the same in each nation. Further, we fit parameters for inheritance or persistence in each class, but not for any cross-national differences in association between classes or in inheritance. This model yields a likelihood-ratio test statistic lower than under the CASMIN model ($L^2=3{,}183$), but it uses 25 more degrees of freedom than the CASMIN model and thus has a slightly larger $bic=-5{,}397$. That is, one would do about as well in fitting the CASMIN data with a best-fitting linear-by-linear association term as with the highly tailored CASMIN model, provided one is willing to admit a modest departure from quasi-symmetry.[21]

Model 7 is similar to model 6, except it constrains the scale values to be equal for origin and destination classes. Its fit is significantly worse than that of model 6 ($\Delta L^2=316$; $df=10$; $bic=210$). However, based upon our earlier comparison of the row and column scores (Figure 2), we prefer model 7 to model 6. That is, we do not think the three reversals in class position between models 6 and 7 are sufficiently large or plausible to warrant a preference for the less constrained model. Note, however, that model 7 fits the data less well than the CASMIN model by all measures except the index of dissimilarity.

Models 8, 9, and 10 depart from the previous specifications by specifying cross-national variation in key parameters of the model. In model 8, we permit the overall level of class inheritance to vary across nations; in model 9, we permit the overall level of linear-by-linear association to vary across nations; and in model 10, we permit both of these forms of cross-national variation. Each of these additions to model 7 is nominally statistically significant, but none of them improves fit when *bic* is the criterion. That is, there are statistically significant improvements in fit relative to model 7 or to model 9 when we fit cross-national variations in class inheritance, and there are statistically significant improvements in fit relative to model 7 or to model 8 when we fit cross-national variations in linear-by-linear association. Yet each of these changes in the model reduces the likelihood-ratio test statistic relative to its degrees of freedom.

Conditional on the equal, homogeneous row and column effects, there is little evidence of cross-national variation, either in class persistence or in linear-by-linear association. Following Yamaguchi (1987), we believe that these tests of cross-national variation are stronger, both statistically and theoretically, than the models on which they are based. That is, each test requires only one degree of freedom for each nation, and they are based on the simple and theoretically appealing concepts of greater or lesser class inheritance and greater or lesser rank mobility.

Model 11 is a baseline for our addition of terms for linear association in measured variables and symmetric sectoral effects. It simply fits cross-nationally invariant terms for class persistence. Clearly, class immobility in itself accounts for a substantial share of the association in the 12×12 tables for Western Europe; the model accounts for more than half of the L^2 under conditional independence. Model 12 fits the parameters for class inheritance, plus two linear-by-linear interaction terms, one for occupational socio-economic status (SES), and the other for occupational prestige (PRES). The SES scores are from the international scale of socio-economic status proposed by Ganzeboom *et al.* (1992), except that we have interpolated scores for classes III*a* and III*b* and for classes IV*c* and IV*d* using Hout's (1989) socio-economic scores for Irish occupations.[22] Prestige is from the Hope–Goldthorpe scale (Goldthorpe and Hope, 1974). These two terms are highly significant, separately and collectively; under model 12, $L^2 = 5{,}021$ with $df = 833$, and $bic = -3{,}760$.

In model 13 we add three terms for occupational sector: SE2, which pertains to any combination of origins and destinations in classes I, IV*a*, IV*b*, IV*c*, and IV*d*; AG, which pertains to any combination of membership in classes IV*c*, IV*d*, and VII*b*; and AG.SE2, which pertains to any combination of origins and destinations in the intersection of SE2 and AG, that is, classes IV*c* and IV*d*. Erikson and Goldthorpe (1987*a*, 1987*b*) and Hout (1984*a*, 1989) find that categories of self-employment show special affinities and that mobility within the agricultural sector is higher than might otherwise be expected; we have extended their specifications by adding a parameter for the combination of self-employment and the agricultural sector.

This model fits better than any we have considered, except constant quasi-symmetric association ($L^2 = 2{,}455$; $df = 830$; $bic = -6{,}294$), and it is superior in fit both to the CASMIN model and to the association models. While the model of constant social fluidity has a lower L^2, model 1 uses so many more degrees of freedom that its *bic* statistic is somewhat higher than that of model 13. Although we have added the sectoral distinctions to the model after socio-economic status and prestige, they obviously lead to a very large reduction in the L^2 (compare lines 12 and 13 of Table 4). A model with the diagonal and sectoral effects, but no effects of hierarchy, yields $L^2 = 5{,}303$ with $df = 832$, and $bic = -3{,}468$. Thus, if we put the sectoral variables into the model first, and then add the hierarchical effects, socio-economic status and prestige lead to a very large reduction in the test statistic. Both the sectoral and the hierarchical effects are highly significant, and both contribute substantially to the good fit of model 13. For this reason, we believe that if Featherman *et al.* (1975: 357) overstated the case for occupational socio-economic status as the central dimension of social mobility, then Erikson and Goldthorpe (1987*a*:72) equally overstate the case for the preponderance of sectoral effects. We think that both are highly significant in the CASMIN data.

Model 13 is our preferred specification of the 12×12 mobility classifications for Western European nations, and we use it as a baseline in a provisional effort to identify cross-national variation in mobility regimes. This model of cross-nationally constant mobility contains terms for the nation-specific distributions of origin and destinations and seventeen terms for the association between origins and destinations: twelve main effects of class immobility, and one effect for each of self-employment, agricultural sector, the interaction effect of self-employment and the agricultural sector, prestige, and socio-economic status. Excepting a few of the diagonal effects, the parameters of model 13 are large and highly significant. The effect (with standard error in parentheses) of the self-employed sector is 0.453 (0.034); that of the agricultural sector is 1.730 (0.070); and the effect of self-employment within agriculture is 1.635 (0.090). One might think of each of these coefficients as an immobility effect that pertains, not to a single diagonal cell in the classification, but to a block of cells. Thus, taking account of other effects in the model and regardless of class, persistent self-employment is $e^{0.453} = 1.57$ times as likely as any combination of non-self-employment with any other sectoral location.

The effect of socio-economic status ($\times 1000$) is 1.745 (0.056), and that of prestige ($\times 1000$) is 0.371 (0.056). That is, the odds on a higher status or prestige destination increase with rising prestige or status of origin. From these coefficients we can calculate the average effect of a one-point increase in status and prestige of origin on the log-odds of one destination versus another. Suppose we wish to see the effect of origin status on the log-odds of a destination in the professional and managerial lower grade category (class II) versus a destination in a semi-skilled or unskilled manual occupation (class VII*a*). There is a socio-economic status gap of 27 points between destination classes II and VII*a*, so the effect of status on logit (II:VII*a*) is $0.001745 \times 27 = 0.047$. The prestige gap for these origins is 38 points, so the effect of prestige on logit(II:VII*a*) is $0.000371 \times 38 = 0.014$. On that basis the model predicts that sons of routine non-manual employees (SES = 57; prestige = 42) will have an advantage of 1.10 on the logit scale compared with sons of skilled manual workers (SES = 35; prestige = 38) and a disadvantage of the same magnitude (1.10) compared with the sons of upper-grade professionals and managers (SES = 71; prestige = 73). Translating from the logit scale to probabilities, we note that if the marginal parameters and the effects of other variables in the model make the expected probability of a destination in class II instead of class VII*a* equal 0.50 (i.e. the two destinations are equally likely) for sons of routine non-manual employees, then status and prestige together will imply a probability of 0.25 for sons of skilled manual workers and 0.75 for sons of upper-grade professionals and managers.[23]

Before turning to cross-national variations in mobility, we consider the consequences of specifying the hierarchical dimension in the way that Erikson and Goldthorpe do, i.e. with a social-distance term for the number of hierarchical levels crossed. In model 14, we substitute the two vertical mobility contrasts of the CASMIN model, HI1 and HI2, for the linear effects of prestige and socio-economic status. While model 14 fits substantially better than the CASMIN model, it is substantially worse than model 13. The linear-by-linear specification of vertical mobility is superior to that of the CASMIN model. The main contribution of the CASMIN specification would appear to be in drawing attention to sectoral effects, which we have coded in a somewhat different way, rather than to its specification of vertical mobility or of specific affinities between classes.

We searched for a model of cross-national variation in mobility by adding selected interactions of nation with terms in model 13 and selecting backwards, that is eliminating those sets of interactions that contributed least to the explanation of cross-national differences. In fact, we began by adding all of the interactions of terms in model 13 with nation, except that we introduced only one global effect for class inheritance in each nation. That is, we permitted inheritance to interact with class (as in model 13) and with nation, but not with both class and nation. There were 36 degrees of freedom for cross-national variation in the least restrictive model we considered (model 15), that is, 6 degrees of freedom for nations for each of the six mobility effects: inheritance, agricultural sector, self-employment, self-employment in agriculture, prestige, and socio-economic status. When all of these terms were added to the model, the value of *bic* increased from $-6,294$ to $-6,157$ ($L^2 = 2,212$ with 794 *df*); that is, the decrease in the L^2, $2,455 - 2,212 = 243$, was not sufficient relative to its degrees of freedom. Lest one suspect that model 13 lacks power to detect cross-national variation, we note that when all eight of the effects in the CASMIN model are interacted with nation, the L^2 declines from 3,416 to 3,257 in the same data, an improvement of just 159 on 48 degrees of freedom. That is, the CASMIN model has less power to detect cross-national differences than does model 13.

If we relied upon *bic* as the criterion for including cross-national interactions and wanted to improve the fit of model 13, we would not add any terms to the model. However, because of our strong prior interest in cross-national variation, we deleted just the two weakest sets of cross-national interaction effects, those pertaining to self-employment in agriculture and to self-employment, and then refitted the model. The cross-national interactions of immobility and of agricultural employment were also relatively weak statistically, but we decided to retain them. We think it is of some import that the least significant cross-national interactions in the model pertain to sectoral and inheritance effects. Saving 12 degrees of freedom by dropping cross-national variation in the effects of SE2 and SE2.AG raises L^2 by 40 and produces model 16, for which *bic* is 50 points closer to zero than in model 13 ($L^2 = 2,252$; $df = 806$; $bic = -6,244$). That is, model 13 is still preferable to model 16, and we urge caution in the interpretation of cross-national differences in the parameters of model 16.

The interaction effects in model 16 are reported in Table 6. The diagonal effects are of interest because they measure the extent to which class-specific persistence departs from the persistence implied by the effects of hierarchy and market sectors. While persistence in class I is fitted well by the final model, there is substantial persistence in classes IV*a*, IV*b*, and IV*c*, above and beyond that implied by the sectoral effects. On the other hand, net of sector effects, there is net movement out of class IV*d*. There is also substantial immobility in classes II, III*b*, V, and VI, but not at levels comparable to the immobile self-employed or agricultural classes. The nation × diagonal interactions show that only in two nations, Ireland and Northern Ireland, are there distinctive overall levels of class immobility, which are each about 35 to 40 per cent higher than in other nations. The nation × agriculture interactions show that persistence in the agricultural sector is distinctively high in England and Wales and in Scotland and distinctively low in Sweden.

The effects of prestige and of socio-economic status vary substantially. There is some indication that the variations are inverse to one another. That is, where the effect of socio-economic status is large and positive, the effect of prestige is smaller and, in some cases, negative. Thus, France and England and Wales have relatively large, positive effects of prestige and lesser effects of socio-economic status, while West Germany, Ireland, and Sweden have large, positive effects of socio-economic status and negative effects of prestige. The simple correlation between the SES coefficients and the prestige coefficients is $-.76$, and we have tested whether the effects of these two principles of stratification are truly inverse. This pattern of effects could be an artefact of multi-collinearity, but we do not think that it is.

In this context, there are three logical alternatives to the specification of model 16: (1) the effects of socio-economic status vary across nations, but those of prestige do not; (2) the effects of prestige vary across nations, but those of socio-economic status do not; (3) the effects both of socio-economic status and prestige vary

TABLE 6 *Main effects and nation-specific effects in a model for 12×12×7 tables: Men 25–64 years old, Western European nations in the CASMIN data (N=37 849)*

Term	England and Wales	France	West Germany	Ireland	Northern Ireland	Scotland	Sweden
SES	1.92	1.42	2.38	2.51	1.64	2.07	2.48
Prestige	.21	.84	.14	−.47	−.12	.18	−.56
Self-employed (SE)[a]	.45	.45	.45	.45	.45	.45	.45
Agriculture	2.24	1.66	1.18	1.31	1.72	2.47	.93
SE × Agriculture[a]	1.69	1.69	1.69	1.69	1.69	1.69	1.69
Diagonal[a]							
I	.00	.00	.00	.00	.00	.00	.00
II	.32	.32	.32	.32	.32	.32	.32
III*a*	−.01	−.01	−.01	−.01	−.01	−.01	−.01
III*b*	.24	.24	.24	.24	.24	.24	.24
IV*a*	.27	.27	.27	.27	.27	.27	.27
IV*b*	.81	.81	.81	.81	.81	.81	.81
IV*c*	.99	.99	.99	.99	.99	.99	.99
IV*d*	−.26	−.26	−.26	−.26	−.26	−.26	−.26
V	.30	.30	.30	.30	.30	.30	.30
VI	.41	.41	.41	.41	.41	.41	.41
VII*a*	−.11	−.11	−.11	−.11	−.11	−.11	−.11
VII*b*	.16	.16	.16	.16	.16	.16	.16
Diagonal × Nation	.00	.02	.12	.31	.35	.06	−.01

Note: [a]Constrained to be invariant across nations.

across nations, but they vary directly, not inversely. We tested the first two hypotheses by selecting backwards from model 16, that is, dropping the cross-national variation in PRES or dropping the cross-national variation in SES. Relative to model 16, when we drop PRES.C, L^2 increases by 96 with 6 *df* and *bic* increases by 33; when we drop SES.C, L^2 increases by 64 with 6 *df*, and *bic* increases negligibly. To test the alternative of direct cross-national variation between the effects of PRES and SES, we first fitted a model in which the significant cross-national variations in diagonal and sectoral effects were retained, but PRES and SES had invariant effects. Then, we used the coefficients of PRES and SES in this model to construct a new variable, defined as the combined linear effect of PRES and SES in this model, $\text{COMB} = 0.00036902\ \text{PRES} + 0.0017496\ \text{SES}$. Finally, we estimated a model in which the effects of COMB, but not of PRES and SES, varied across nations. Relative to model 16, L^2 increases by 98 with 6 *df*, and *bic* increases by 34, so a model of inverse variation in the effects of SES and PRES is preferable to that of concomitant variation. Thus, we clearly reject the first and third of the alternatives, while the verdict on the second is less certain, but also tends toward rejection.

The effect of status is much larger than that of prestige, so the overall effect of origins on destinations is positive, i.e. advantaged origins produce advantaged destinations. To illustrate this point, Figure 5 plots the expected logit for a destination in class II versus class VII*a* by origin and nation (standardized so that origin VII*b* has an expected logit of zero as a way of eliminating the national differences in occupational composition). The relationship between origin status and the logit of class II versus class VII*a* destination would be linear if the prestige and diagonal effects were zero. The cross-national differences in the overall steepness of the relationship between origins and destinations reflect cross-national differences in all three effects: status, prestige, and diagonal effects. To summarize the cross-national differences in a single index, with due attention

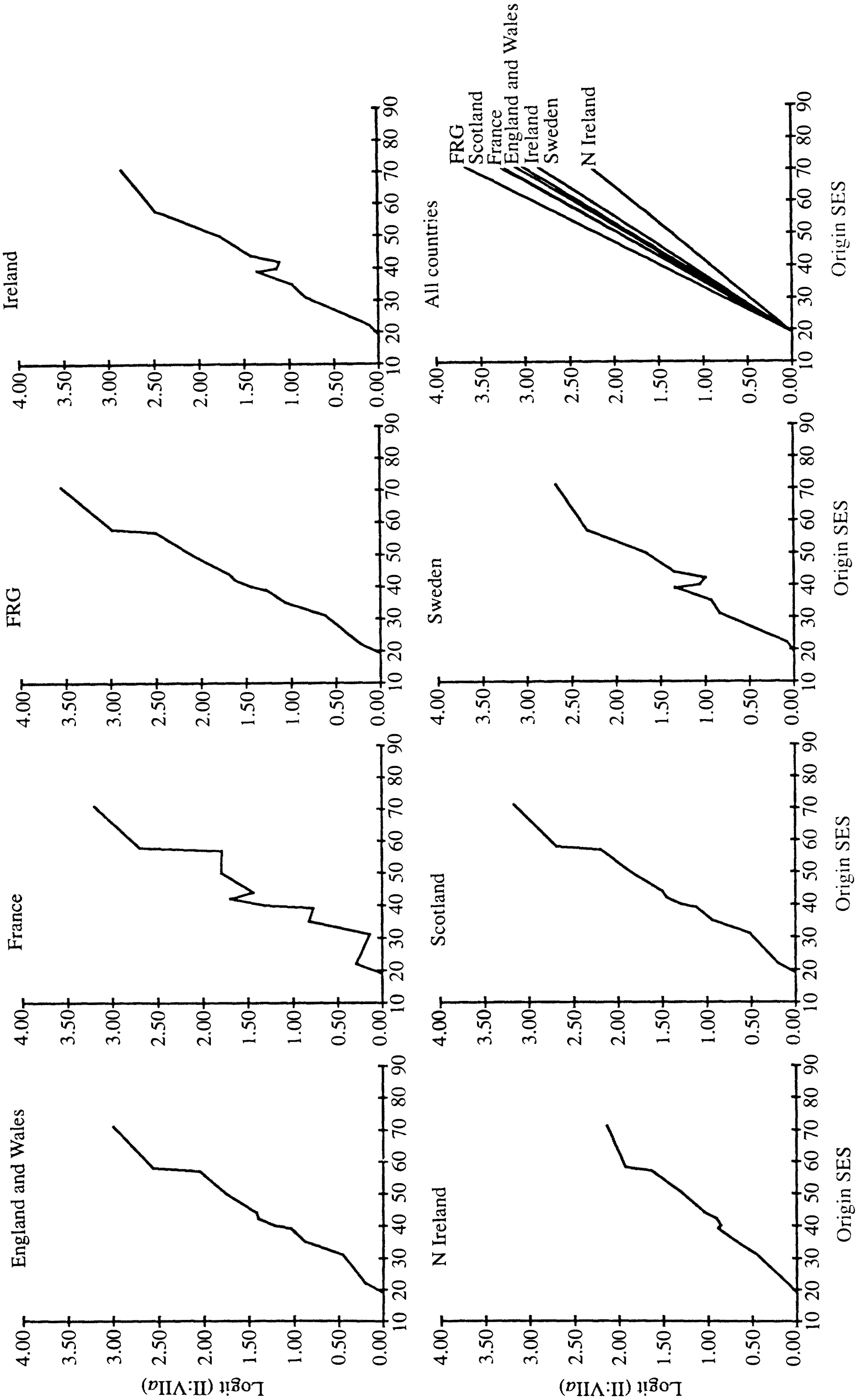

FIGURE 5 *Expected logits under the preferred model: men 25–64 years old, CASMIN data.*

to the information lost in doing so, we have added an eighth panel that smoothes each nation's curve into a straight line (using OLS). This operation shows West Germany to be the most class-stratified of the nations considered, Northern Ireland the least, and the other five clustered at a narrow range of slopes. The range of 'slopes' is from .0437 for Northern Ireland to .0705 for West Germany; the West German value exceeds the Northern Ireland value by 61 per cent.

The parameter estimates in Table 6 suggest one final observation about the methods of the CASMIN study. Rather than carrying out their comparative analysis entirely within the framework of models and methods for the analysis of cross-classified counts, Erikson and Goldthorpe (1987*a*) first estimate their final model with the British and French data and then—claiming that those two mobility regimes are 'central' among the industrial nations—use the expected frequencies for those nations as their model and standard in further cross-national comparisons. Given the fact that several parameters of the final model are cross-nationally invariant, one might take findings in any of the seven nations as a standard. However, among the values of some parameters that differ across nations, we do not see England and Wales or France in a central position. Both cases are similar to Scotland and Sweden in sharing a lower level of immobility than West Germany, Ireland, or Northern Ireland. The holding effect of the agricultural sector in England and Wales is second only to Scotland. England and Wales and France have the two largest positive effects of prestige and two of the smallest effects of socio-economic status. Thus, our findings leave room for doubt about the centrality of the mobility regimes in England and Wales and in France, relative to the other five industrial nations in the analysis.

CONCLUSIONS

The CASMIN model of class mobility proposed by Erikson and Goldthorpe advances our understanding of cross-national differences in social mobility in a number of important ways, most notably by pointing out the extent to which differences in the association between social origins and destinations reflect consequences of public policies that enhance or restrict opportunities. We have respecified the CASMIN model in ways that clarify those cross-national differences. In particular, the problems with the CASMIN model are its application to aggregated occupational categories, its asymmetry, and its treatment of hierarchical or vertical differences among social classes. Our alternative specification uses greater occupational detail, imposes symmetry on the association between origins and destinations, and incorporates a linear-by-linear expression for the effect of origins on destinations similar to that found in regression models.

Taking advantage of these refinements, we have presented estimates of the effect of structural mobility on overall mobility in the CASMIN nations. We find that the patterns of structural mobility in Europe differ more by occupation than by nation. In each nation we find a growing white-collar sector, stable self-employed and blue-collar sectors, and a shrinking agricultural sector. Hungary stands out as the nation with the greatest structural mobility; England and Wales and Scotland have the least.

We conclude that the CASMIN model understates the importance of hierarchy relative to sector and inheritance in the determination of mobility patterns generally as well as cross-national differences. It aggregates classes that are heterogeneous in prestige and socio-economic status; it mis-specifies the vertical position of class aggregates; and it adopts a weak specification of vertical effects. We do not think that it is defensible or necessary to make a case for non-vertical effects on mobility by down-playing the significance of vertical mobility. Our linear-by-linear specification of the hierarchy effect shows a stronger effect of origin prestige and status on destination prestige and status than the CASMIN model shows for the effect of distance between origin and destination on the rate of mobility. Our comparison of the two kinds of models (social distance and linear-by-linear) shows why that would be the case. At the same time, our preferred models also reveal large effects of sectoral location that are not well specified by vertical mobility, even when the terms in vertical mobility are freely scaled.[24]

The principal concern of the CASMIN project is, after all, the *Comparative Analysis of Social Mobility in Industrial Nations*. They and we have assessed their findings against the background of the much-discussed (and long-obsolete) FJH (Featherman, Jones, and Hauser) hypothesis that, in modern nations, at least, the association between origins and destinations is characterized by a 'basic' similarity (Featherman *et al.*, 1975). Erikson and Goldthorpe conclude that the CASMIN model supports the verbal formulation of the FJH hypothesis, because it reveals variations in the degree of association between origins and destinations, but not variations in the kind of mobility processes. That conclusion might be assailed by a sceptical critic because some of their national variants consist of modifications to the design matrices of the CASMIN model, as well as variations in the parameters of the design matrix of the core model. Featherman and colleagues argued for identity across nations in almost all odds ratios between occupational origins and destinations, excepting some unique features of farm inheritance in the American and Australian data that they analysed. Thus, our sceptical critic might claim that Erikson and Goldthorpe have stretched the FJH hypothesis beyond recognition; if variation in parameters and in design matrices are not sufficient to indicate the failure of the FJH hypothesis, what evidence would support that conclusion? In any event, our model applies identically to all nations. In this way our results strengthen the case for basic similarity. We obtain an acceptable fit for all nations with the same parameterization.

Despite the greater parsimony of our model, we do not wish to lean too heavily on the notion of basic similarity. The variations of degree that we find are substantial. Our parameter estimates indicate not only strong main effects of origin prestige and status on destination prestige and status, but also large variations in the sizes of those effects. The point at which a difference of degree overturns a hypothesis of basic similarity is not altogether clear, but we think that future theory and research on cross-national differences in mobility ought to take very seriously the ranges of prestige and status effects and the position of particular nations within those ranges.

NOTES

1. The categories, with the combination of Roman numerals and letters that commonly identify them are: I = professionals and managers, upper grade; II = professionals and managers, lower grade; III*a* = clerical and other routine non-manual workers; III*b* = sales and service workers; IV*a* = proprietors with employees; IV*b* = proprietors without employees; IV*c* = farmers with employees; IV*d* = farmers without employees; V = technicians and foremen; VI = skilled manual workers; VII*a* = semi-skilled and unskilled manual workers; VII*b* = farm labourers. In most of their work, Erikson and Goldthorpe combine I with II, III*a* with III*b*, IV*a* with IV*b*, IV*c* with IV*d*, and V with VI.
2. Throughout the rest of the paper we use Erikson and Goldthorpe's word 'hierarchical' to describe the effect of the high-to-low dimension of social origins. The word 'vertical' is probably more descriptive than 'hierarchical' of our conception of socio-economic and prestige effects (see Hope, 1982; Hauser 1984*a*, 1984*b*), considering the implication of formal authority structure in the word 'hierarchy'. None the less, because we raise so many technical issues in this paper, we chose to leave the semantic issue alone, except for this brief note.
3. In our use of the term 'structural mobility', we adhere to the convention of treating origins as the characteristics of respondents, not their fathers. We firmly reject any notion that the origin distribution is indicative of any point in the past.
4. The employer information was not available in the Hungarian and Polish contexts. Presumably fathers could have had employees if the father operated his shop or farm prior to the installation of communism in these states, but the coders in these studies did not distinguish employing proprietors and farmers in either the origin or destination categories.
5. Another useful way of viewing this test is to think of it as fitting a topological model to the full table in which each level is the collection of cells that make up the individual cells of the CASMIN tables.
6. For example, if workers with origins in class V have an advantage over their counterparts from origins in class VI in moving into upper-grade vs. lower-grade professional and managerial occupations (class I vs. II), then that component of the association between origins and destinations will be reflected in this ratio.
7. These asymmetries use four (not two) degrees of freedom because the asymmetrical treatment of farm origins and destinations implies an asymmetry at each of the three hierarchical levels.
8. This is an apt comparison because the entries in diagonal cells do not affect the test statistics for quasi-independence or for quasi-symmetry.
9. In addition, we specified a distinct parameter for inheritance of each occupational class, but no differences among countries in class inheritance. We estimated this model in GLIM, using our modified version of a macro prepared by Raymond Sin-Kwok Wong.

10. We fitted this model using our modified version of a GLIM macro written by Herbert L. Smith.
11. Employers were not identified in the Hungarian and Polish data-sets, so structural mobility is undefined for farmers and proprietors with employees in these two countries. The undefined values are graphed as if they were zero.
12. These values are low because few men in England and Wales and in Scotland have farm origins, not because of the persistence of farming.
13. Hungary will show an above average positive shift in several categories to counterbalance its extraordinary negative shift in the farmer category.
14. The index of overall redistribution is the standard deviation of the $\log(\alpha_j)$ across categories, i.e. $\sigma_\alpha = [\Sigma_j \log(\alpha_j)^2/J]^{1/2}$, where J is the number of categories.
15. Together these constraints that all the parameters are positive ensure that $F_{ij} > 0$, i.e. that there are no zero or negative expected frequencies and that $\log \mu$, $\log \alpha_j$, $\log \beta_j$, and $\log \gamma$ all exist.
16. The data in Figure 4 are the logits expected under the social-distance and linear-by-linear models for the bivariate normal tabulation in Goodman (1984, Ch. 4, Table 1). We chose these data because they were a convenient source of uncomplicated data with a strong association between origins and destinations. They reveal nothing about social mobility, but they highlight this important difference between mobility models.
17. The two broad groupings are those origins for which the better destination, j in the pair (j, j'), is less than or equal to the origin on the X-scale (i.e. $X_j \leqslant X_i$), compared with those origins for which the better destination is greater than the origin on the X-scale (i.e. $X_j > X_i$).
18. The tendency for frequencies in mobility tables to cluster on the diagonal due to immobility is not well modelled by the linear-by-linear model, so diagonal terms are added to the model: either a dummy variable for each diagonal cell or a set of more general terms as in Hout (1984*a*) or Erikson and Goldthorpe (1987*a*, 1987*b*). Diagonal clustering is an intrinsic feature of the social-distance model because all diagonal cells have a distance score of zero.
19. The relationship between origin and downward mobility under the linear-by-linear model is indeterminate, but it will be positive under most values of association parameters (Hout, 1984*b*).
20. The *bic* for the CASMIN model is actually more favourable in the replication than in the original data. Although this is a surprise, it is related to the low power of the replication. Even the model of independence fit to the replication data has a more favourable *bic* ($L^2 = 230.99$; $df = 81$; $bic = -292$) than the CASMIN model has in the original data.
21. This model is not nested under quasi-symmetry because the distances between origin classes are not necessarily the same as the distances between destination classes.
22. We began with the scores reported for 10 of the CASMIN classes by Ganzeboom *et al.* (1992). We interpolated scores for classes III*a*, III*b*, IV*c*, and IV*d* by an iterative procedure, regressing the Ganzeboom *et al.* SES scores on Irish SES scores (Hout, 1989: 69) and replacing the observed (or previous) values of the international SES scores for the four affected classes by the predicted values at each iteration until the process converged on the scores: I = 71; II = 58; III*a* = 57; III*b* = 39; IV*a* = 50; IV*b* = 40; IV*c* = 42; IV*d* = 22; V = 44; VI = 35; VII*a* = 31; VII*b* = 19.
23. The symmetry of this calculation is a function of our choice of 0.50 as the baseline probability. Baselines greater than 0.50 will result in greater differences in expected probabilities between class III*a* and class VI than between class III*a* and class I, while baselines less than 0.50 will result in the larger differences between III*a* and I than between III*a* and VI.
24. That is, model 13 is superior in fit to the association models (6 and 7), as well as to the CASMIN model.

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