LEAST-SQUARES ESTIMATION OF EFFECTS ON INTER- AND INTRAGENERATIONAL OCCUPATIONAL TRANSITION PROBABILITIES[1]

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Introduction

A great deal of attention has been devoted recently to the development of models for the estimation of the relative influence of a set of explanatory conditions on binary dependent variables [8,9,10,11,12,13,15,18]. The present paper presents the application of one of the suggested techniques to two social processes which lend themselves to conceptualization within this framework: intergenerational and intragenerational occupational mobility.

Individuals can be characterized as having certain probabilities of making a transition from a prescribed set of "origin" positions into some set of "destination" positions. We will utilize a model which provides for the introduction of independent conditions to explain variations in the probabilities of movement out of these origin positions.

Formulation and Estimation

The simplest approach to the problem is to employ an ordinary least-squares (OLS) model

$$Y_i = \beta_0 + \sum_1^j \beta_j X_{ij} + \varepsilon_i \qquad (1)$$

in which Y_i assumes a value of unity if the i^{th} person was a mover from a particular origin position to a specified destination position during the time period of interest [cf. 12,16]. Since the expected value of Y_i is the population transition probability between the origin and destination

$$E(Y_i) = Pr(Y_i = 1) = P_i$$

positions, equation (1) can be expressed as the linear probability function (LPF):

$$E(Y_i | \sum_1^j X_{ij}) = Pr(P_i | \sum_1^j X_{ij}) = \beta_0 + \sum_1^j \beta_j X_{ij} \qquad (2)$$

The unstandardized regression coefficients of the LPF are interpreted straightforwardly as the relative contribution of each of the independent variables to the transition probability [16].

As pointed out by Theil [18] and Neter and Maynes [13], among others, this model has two major defects. The first difficulty involves the distribution of the error variances. It can be shown that the error variances have a binomial distribution [7,8],

$$V(\varepsilon_i) = P_i(1 - P_i) = P_i Q_i$$

thus the homoscedasticity assumption of OLS can not be met. This problem is readily solved by employing a weighted least-squares (WLS) solution for the parameter estimates [7,8,15,18]. The WLS estimators are minimum variance linear unbiased estimators of the LPF.

The second problem with the LPF is that the expected value of the dependent variable, the predicted transition probability, can exceed unity or be less than zero for any given observation. This undesirable outcome is the result of the left-hand side of equation (2) being constrained to vary between 0 and 1 while the right-hand side is under no such limitation.

Two recommendations have been advanced for coping with this problem. Speare [15] and Huang [7] suggest that the simplest solution might be to truncate the predicted transition probability at 0 and 1. However, since the first derivative of the conditional expectation is no longer continuous, an exact solution for the parameter estimates is not possible and an iterative estimation procedure must be employed [15]. While this approach maintains the integrity of the interpretation of the unstandardized coefficients, it is costly in computational terms.

Theil [18] and Goodman [9,10], among others, recommend a second strategy: transform the dependent variable such that the 0,1 constraint no longer presents difficulties. Specifically, Theil suggests using the "logit" (logarithmic unit) transformation:

$$L_i = \log_e \frac{P_i}{Q_i}$$

which has the effect of producing a dependent variable, L_i, which approaches $+\infty$ as the probability of moving, P_i, approaches unity and $-\infty$ as P_i approaches zero. The statistical model then becomes:

$$E(L_i | \sum_1^j X_{ij}) = \beta_0 + \sum_1^j \beta_j X_{ij} \qquad (3)$$

Estimates of the parameters then measure the relative effects of the explanatory variables on the logit itself. After solving for the WLS estimates of the parameters, the expected transition probabilities can be readily retreived:

$$\hat{P}_i = (1 + e^{-\hat{L}_i})^{-1}$$

By comparing directly estimates derived from the LPF, the truncated LPF, and the logit models, Speare [15] concludes that the latter two are superior since they tend to assign larger effects to the independent variables thus allowing for the incorporation of a greater number of explanatory conditions in a given model. While the truncated LPF does maintain one advantage over the logit specification, the clear interpretation of the unstandardized coefficients, it is our position that the logit model is to be preferred on theoretical grounds as well as for its greater computational facility.

The truncated LPF implies that it is possible that the predicted transition probability for any given observation *may* be 0 or 1; it is difficult for us to believe that the probability of mobility would ever reach these extremes. It is more likely that there is always a small probability of movement out of an origin position, although this probability may be quite small. On the other hand, it is equally difficult for us to conceive of a situation in which it is certain that an individual will move regardless of the circumstances. The logit specification implies that the probability for stayers never attains zero, and likewise, the probability for movers never reaches unity. We find these implications more satisfying than those implied by the truncated LPF. The logit specification will be used in this paper.

City Effects on Intergenerational Mobility

The first application of the model is to an exploration of the extent to which the social context of urban centers influences intergenerational occupational mobility. Students of social mobility have increasingly turned to the city as the locus of change in the structure of stratification since it is the urban milieu which tends to generate the key processes having a direct effect on mobility rates (see, for example, Schnore [14]). In brief, we argue, as has Duncan [5], that American cities are organized in a system in which functional specialization is overlaid with a set of dominance relationships based upon the relative concentration of financial and commercial activities in the city.[2] Regardless of functional specialization, as indexed by industrial composition, it is hypothesized that a more dominant place will incorporate a larger proportion of highly ranked coordinative and administrative positions. In addition, it is expected that such city effects will be more pronounced for those individuals who have remained in their place of origin, since only then can the differential opportunity structure be expected to operate at the individual's point of entry into the labor market. The introduction of a control for migration status is thus essential to establish the existence of such an effect.

The data to be utilized in the following analysis are those collected by the Bureau of the Census for Blau and Duncan's [2] major survey of the extent and sources of social mobility in the United States, "Occupational Changes in a Generation" (OCG). As an adjunct to the Bureau's monthly "Current Population Survey" (CPS), the OCG questionnaire, or a follow-up interview, was administered to a national sample of 20,700 males between 20 and 64 years of age.[3] Since the present analysis will utilize SMSA's as units of analysis, it is crucial that the sample be representative of intra-SMSA populations. Our best assurance of this is that in the 1962 CPS all SMSA's were coterminous with primary sampling units. Only the fifteen largest SMSA's for which OCG data were collected were coded separately, and it is to these that the analysis must be restricted.

The dependent variable is the observed probability of mobility from any origin status (father's occupation) upward into a destination status (son's occupation) where coordinative or administrative functions are performed; coordinative or administrative positions are operationalized here as professional, managerial, and clerical positions. This probability is calculated from the relative frequency of movers in categories determined by the independent variables:

$$X_1 = \begin{cases} 1 & \text{If the person resides in the same city he did at age 16} \\ 0 & \text{Otherwise} \end{cases}$$

TABLE 1
UPWARD MOBILITY INTO COORDINATIVE POSITIONS DETERMINED BY POSITION OF CITY OF RESIDENCE ON URBAN SYSTEM DIMENSIONS AND MIGRATION

Position on Industry Dimension	Migrants		Non-migrants	
	Position on Metropolitan Functions Dimension			
	High	Low	High	Low
	Probability of Upward Mobility			
Manufacturing	0.301 (923)*	0.320 (263)	0.287 (487)	0.234 (286)
Service	0.323 (646)	0.378 (130)	0.302 (463)	0.279 (53)
	Estimated Probabilities			
Manufacturing	0.305	0.297	0.267	0.260
Service	0.336	0.327	0.296	0.288
	Discrepancies**			
Manufacturing	-0.004	0.023	0.020	-0.026
Service	-0.013	0.046	0.006	-0.009

*Number of cases in cell. For purposes of inference, these N's have been adjusted to approximate simple random sampling.

**$\chi^2 = 4.614$ (NS)

$$X_2 = \begin{cases} 1 & \text{If the person resides in a city high on the metropolitan functions hierarchy} \\ 0 & \text{Otherwise} \end{cases}$$

$$X_3 = \begin{cases} 1 & \text{If the person resides in a manufacturing-oriented city} \\ 0 & \text{Otherwise} \end{cases}$$

The first panel of Table 1 indicates the observed values for the conditional probabilities of upward mobility into coordinative positions, along with the corresponding adjusted sample sizes.

Defining p_i as the observed probability of occupational mobility, in the i^{th} subset, the observed logits were obtained:

$$L_i^* = \log \frac{P_i}{1 - p_i} \qquad (5)$$

The following linear logit specification was employed:

$$\hat{L}_i^* = \hat{\beta}_0 + \hat{\beta}_1 X_1 + \hat{\beta}_2 X_2 + \hat{\beta}_3 X_3 \qquad (6)$$

Table 2 presents the WLS estimates of these coefficients and their asymptotic standard errors. As expected, migrants are significantly more likely to be upwardly mobile into coordinative positions, regardless of city of residence.[4] Our initial hypothesis, that residence

TABLE 2

WLS ESTIMATES OF THE MAIN EFFECTS OF EXPLANATORY CONDITIONS ON UPWARD MOBILITY INTO COORDINATIVE POSITIONS

Effect	Estimated Coefficient	Standard Error
Constant	-0.882	0.063
Due to non-migration	-0.185	0.077
Due to residence in a city low on metropolitan hierarchy	-0.041	0.088
Due to residence in a service-oriented city	0.142	0.078

in a city ranked high on the metropolitan functions hierarchy will enhance the probability of upward mobility into coordinative positions, is unconfirmed, since this estimate is exceeded in size by its standard error. It is interesting that a small, but nonsignificant interaction between migrant status and rank of city on the metropolitan hierarchy is evident in the discrepancies between predicted and observed probabilities presented in the third panel of Table 1.[5] Here, the relationship is in the hypothesized direction for non-migrants, but not for migrants.

An unanticipated finding is the marginally significant effect of residence in service-oriented cities evident in Table 2. It is our presumption that such an effect is explainable by variations in the occupational composition of manufacturing vs. service-oriented cities: places with a relatively high concentration of service functions simply contain a higher proportion of coordinative positions to be filled. To test this hypothesis we apply a procedure developed by Deming [3] and applied to mobility tables by Duncan [4] in which the entire table for each city is proportionally adjusted to a new set of marginals: the distributions of sons by their own occupations and by their fathers' occupations. In the present case these adjustments are to the marginals for all SMSA's with population over one million. The question this procedure enables us to answer is whether or not the observed effect of residence in a service-oriented city is produced by an opportunity structure possessing a greater demand for individuals to fill coordinative positions, i.e. will the proportional adjustment of the mobility matrices for individual SMSA's allow us to account for this effect, or are there some additional unmeasured variables which operate to produce nonproportional differences across the fifteen matrices.

As the first panel of Table 3 attests, the proportional adjustment of the mobility matrices rather drastically reduces variability among the observed probabilities. In Table 4 we can see that the previously observed relationship between the probability of upward mobility and residence in service-oriented cities disappears, indeed is reversed to a non-significant degree. Apparently the observed effect of residence in service-oriented cities is wholly explicable by differentials in the opportunity structures of such cities, as opposed to those of manufacturing cities. Again, the model contains no significant interactions.

Industrial Development and Intragenerational Occupational Movers and Stayers

The second example comes from a comparative investigation of the short-term effects of the introduction of a large-scale highly automated steel plant in an agrarian region of the Midwest in 1966.[6] The basic research problem is to assess the relative influence of several explanatory variables on the probability of a given worker changing position in the occupational structure during the period of industrial development, 1966 to 1971. In short, we would like to specify the conditions which affect the transition rates among occupations. Our major interest here is not upon transition rates between specific origin and destination occupations, but upon differentiating occupational movers from stayers. Data for the analysis came from retrospective work histories collected in 1971 from a sample of 778 employed heads of households in two areas of Illinois. The first area comprised townships surrounding the plant location and is designated as the "experimental" region; the second area was a "control" region that has not undergone industrial development.

TABLE 3

OBSERVED, PREDICTED, AND DISCREPANCIES IN PROBABILITY OF UPWARD MOBILITY BY POSITION OF CITY OF RESIDENCE ON URBAN SYSTEM DIMENSIONS AND MIGRATION, OCCUPATIONAL DISTRIBUTIONS ADJUSTED

Position on Industry Dimension	Migrants		Non-migrants	
	Position on Metropolitan Functions Dimension			
	High	Low	High	Low
	Probability of Upward Mobility			
Manufacturing	0.305 (923)*	0.305 (263)	0.293 (487)	0.299 (486)
Service	0.299 (646)	0.274 (130)	0.281 (463)	0.375 (53)
	Estimated Probabilities			
Manufacturing	0.303	0.307	0.295	0.300
Service	0.296	0.301	0.288	0.293
	Discrepancies**			
Manufacturing	0.002	-0.002	-0.002	-0.001
Service	0.003	-0.027	-0.007	0.082

*Adjusted sample N's.

** $X^2 = 2.344$ (NS)

TABLE 4

WLS ESTIMATES OF MAIN EFFECTS OF EXPLANATORY CONDITIONS ON UPWARD MOBILITY INTO COORDINATIVE POSITIONS, OCCUPATIONAL DISTRIBUTIONS ADJUSTED

Effect	Estimated Coefficient	Standard Error
Constant	-0.834	0.063
Due to non-migration	-0.038	0.077
Due to residence in a city low on metropolitan hierarchy	0.024	0.087
Due to residence in a service-oriented city	-0.033	0.079

Three explanatory variables were used in trying to account for the overall mobility rates: (1) the region in which the worker resided; (2) the length of time the worker has spent active in the labor force; (3) and whether or not the worker was employed in "blue collar" types of employment.[7] The following definitions were used:

$X_1 =$ 1 If the worker resided in the "experimental" region during the period 1966 to 1971

0 Otherwise

$X_2 =$ 1 If the worker had been active in the labor force for more than 10 years as of 1966

0 Otherwise

$X_3 =$ 1 If the worker was employed in a "blue collar" occupation at the beginning of industrial development

0 Otherwise.

After partitioning the data by these three conditions, the proportions of occupational movers were computed for each subclassification. These proportions are presented in the first panel of Table 5.

After transforming these observed probabilities by equation (5), the linear logit model, (6), was estimated using WLS. The weighted least-squares estimates of the coefficients, and their asymptotic standard errors, are presented in Table 6. It was found that the effects of being in the experimental region and having been in the labor force for more than ten years tend to decrease the probability of occupational mobility whereas having a blue collar occupation at the beginning of the industrial development period tended to increase the probability of mobility. The largest influence of the transition probabilities was the length of time spent in the labor force while regional residence tended to have the smallest impact.

The third panel of Table 5 presents the discrepancies between the observed and predicted proportions of occupational movers. It should be noted from Table 5 that the linear logit model predicts fairly well for all the subclassifications except for white collar workers in the control region who have been in the labor force for less than ten years. For this group the model underestimates the observed transition

TABLE 5

OBSERVED, PREDICTED, AND DISCREPANCIES IN PROBABILITY OF OCCUPATIONAL MOBILITY BY REGION, OCCUPATIONAL CATEGORY, AND TIME IN LABOR FORCE

	Experimental Region		Control Region	
	Blue Collar	White Collar	Blue Collar	White Collar
	Probability of Mobility			
In Labor Force Less Than Ten Years	0.346 (78)*	0.250 (36)	0.333 (30)	0.471 (17)
In Labor Force More Than Ten Years	0.247 (291)	0.158 (152)	0.250 (120)	0.241 (54)
	Estimated Probabilities			
In Labor Force Less Than Ten Years	0.348	0.283	0.398	0.328
In Labor Force More Than Ten Years	0.234	0.184	0.274	0.218
	Discrepancies**			
In Labor Force Less Than Ten Years	-0.002	-0.033	-0.065	0.143
In Labor Force More Than Ten Years	0.014	-0.026	-0.024	0.023

*Number of cases in cell.

** X^2 = 2.672 (NS)

TABLE 6

WLS ESTIMATES OF MAIN EFFECTS OF EXPLANATORY CONDITIONS ON INTRAGENERATIONAL MOBILITY

Effect	Estimated Coefficient	Standard Error
Intercept	-0.719	0.238
Due to residence in the experimental region	-0.212	0.181
Due to being in the labor force more than ten years	-0.562	0.192
Due to having a "blue collar" occupation	0.305	0.183

probability rather severely.

Conclusions

In summary, this paper has focused on some of the problems associated with estimating the relative influence of a set of explanatory conditions on binary dependent variables and presented two applications of a linear logit model to the prediction of the probability of movement from some set of origin positions to a set of destination positions. These applications have demonstrated the utility of such a model in introducing independent variables to account for specific patterns of movement in both inter- and intragenerational occupational mobility matrices.

FOOTNOTES

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[2]Applying suggestions by Galli [6] of a multiple variable approach to community structure and Winsborough [20] of a Q-factor analytic technique, we sought to test the expectation of Duncan [5] that a system of large urban places can be described by the two dimensions discussed above and rank the 55 largest SMSA's on them. See Wanner [19] for a more complete treatment of this procedure.

[3]For a fuller discussion of the survey design and characteristics of the data set, see Blau and Duncan [2] .

[4]As Blau and Duncan [2] indicate, the greater achievement among migrants is explainable primarily by their generally superior backgrounds, especially son's education and father's SES.

[5]Theil [18] suggests what is essentially a Chi-square goodness-of-fit test between the observed and predicted probabilities for testing the validity of the model.

[6]For further details concerning this investigation see Beck [1] and Summers [17].

[7]"Blue collar" occupations included farmers, craftsmen, operatives, service workers, and laborers.

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