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A CONDITIONAL LOGIT MODEL OF INTERNAL MIGRATION: VENEZUELAN LIFETIME MIGRATION WITHIN EDUCATIONAL STRATA

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ABSTRACT

The conditional logit framework is used to describe how the mutually exclusive and exhaustive probabilities of locational choice, including not migrating, depend on a set of conditioning variables. Aggregate data from the Venezuelan 1961 Census are used to estimate this model for males within four educational strata. The logit model estimated by conventional linear procedures fits the 400 cell contingency table better than does the double log-probability function associated with the "gravity" model. Interregional wage differences are relatively larger among less educated, who are also the groups that migrate least frequently. More educated men appear more responsive to destination wages and less deterred by distance. Though these differences in migratory behavior by educational class may account for interregional wage variation, other hypotheses also need to be investigated for this common pattern. Rapid regional population growth also influences migration, by deterring entry. Only men with a secondary or higher education notably avoided destinations where unemployment was high. Thus, the Harris-Todaro model of migration and development which assumes urban unemployment governs migration is not consistent with the behavior of the majority of the labor force in Venezuela which has relatively little schooling.

I. INTRODUCTION

Much empirical research on migration and its determinants involves estimating double logarithmic equations of approximately the following form:*

$$(1)\quad \ln m_{ij} = \beta_0 + \beta_1 \ln N_i + \beta_2 \ln N_j + \beta_3 \ln D_{ij} + \sum_{k=4}^{\bar{n}} \beta_k \ln X_{k-3,i} + \sum_{l=n+1}^{2n-3} \beta_\ell \ln X_{\ell-n,j} + e_{ij} \qquad (i,j = 1,\ldots,z\ ;\ i \neq j)$$

where m_{ij} is the gross rate of migration from region i to region j per person, N_i, at risk of migration in region i during a specified time period, N_j is the number of persons in region j, D_{ij} is the distance from region i to region j, X_i and X_j refer to conditioning characteristics of region i and j, and e_{ij} is a disturbance representing errors in measurement, functional form, and numerous, hopefully minor, omitted factors.

Since Cary (1858-59) first reasoned that human mobility behaved according to "laws" of social interaction, many bodies of data have been fit, with some success, to forms of the "gravity model" of migration:

$$(2)\quad M_{ij} = N_i m_{ij} = \frac{N_i^{\beta_1+1} N_j^{\beta_2}}{D_{ij}^{-\beta_3}}\ Z\,(X_{1,i},\ldots X_{n-3,i}, X_{1,j},\ldots,X_{n-3,j}), \qquad (i,j = 1,\ldots,z; i \neq j)$$

in which it is assumed that gross flows of migrants from one location to another are directly proportional to the population at origin and destination, and inversely proportional to the distance between regions, and possibly conditional on a function Z of other attributes or forces, represented by the X's.

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* See, for example, Beals, et. al., 1967; Greenwood, 1969 a,b, 1971 a, b; Levy and Wadycki, 1972 a,b, 1974 a,b.

To estimate a relationship determining migration, gross migrant flows are usually normalized. Statistical criteria for the choice of a dependent variable include that the residual stochastic disturbance in the resulting relationship is independent of explanatory variables to avoid bias, and of known or constant variance to increase efficiency. Dividing equation (2) by origin population, N_i, and taking logarithms of this gross migration rate equation could explain the origins for estimation equation (1).* Rarely are statistical or economic reasons for this or other normalizations of gross migration offered in the literature.**

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* It is often assumed that individual behavior is independent of what others do, at least in the short run, and hence if a stable stochastic process describes the determinants of individual migration probabilities, the parameters to this process might be estimated from migration frequencies within population aggregates. If individual migration probabilities are randomly and normally distributed, the gross migration rate, m_{ij}, for aggregates is an unbiased estimator for the underlying individual probability, P_{ij}. Even in this case, the variance in estimates of m_{ij} will tend to be inversely proportional to N_i. Generalized least squares estimates are then more efficient through the application of appropriate population weights. If logarithmic equation (1) is estimated, weights must be modified accordingly. For appropriate weights with the logit model see Cox (1970, p. 106).

** The lack of normalization is more difficult to understand. Sahota's (1968) study of Brazilian migration relies largely on double log regressions, but considers as his dependent variable gross lifetime migration flows among the Brazailian states as recorded in the 1950 Census. Though he interpretes his findings in terms of individual responsiveness of migration to a host of variables, it is not clear how he can relate his estimates to the micro economic behavioral model he posits. Sahota's specification also contradicts the classical assumption of homoscedastic disturbances in the regression equation, and since the origin population size tends to be correlated with other determinants of migration, bias as well as loss of efficiency occurs (Schultz, 1969) If it were not for Sahota's specification, some parallels might be sought between his study of Brazilian migration circa 1950 and this investigation of Venezuelan migration as of 1961. Greenwood (1969b) also analyzes gross flows in Egypt, without comment or justification.

The double logarithmic probability function (1), in addition to its dubious inspiration from 19th century physics, also neglects information embodied in the frequency of nonmigration, namely the m_{ii}'s, and admits predictions of "probabilities" in excess of one. Clearly, there are statistical reasons for considering other approaches to migration as a multiple choice process; one such approach is explored in this paper: the conditional logit.

Given an improved statistical model of migration, it should be possible to test economic hypotheses that underlie much thinking about the development process. A widely accepted class of models of migration, labor factor market distortions, and development has distinct but untested implications with respect to the responsiveness of migration to wage and employment rates. Those are tested below, given data limitations. Disaggregating migration by educational attainment and sex is helpful in reducing the heterogeneity of the labor force. This simple but uncommon procedure clarifies the functioning of interregional labor markets, and provides an empirical basis for evaluating regional dualism in a low income country.

The plan of this paper is as follows. The characteristics of logit framework are discussed in section III, and adapted to test several economic hypotheses pertaining to migration. Some tabulations of migration data for Venezuela are presented in section IV, and then these aggregate data are used to estimate and compare the explanatory power of the logit and conventional gravity model of migration. Before considering the statistical model, section II reviews some salient issues in migration research and their policy content.

II. Economic Issues in the Study of Migration

Migration as a Supply Response of Labor

Migration has been variously analyzed as a long-term human capital investment (Sjaastad, 1962), as a selective response of more energetic and adaptable individuals to the changing distribution of economic activity (Kuznets, 1964), and as the summation of presumably asymmetric "push and pull" factors associated with the individual and his environment (Lee, 1966). All of these approaches posit an individual moving to where his future appears most attractive, or more formally, to where he maximizes the present (discounted) value of future streams of benefits minus costs (opportunity, direct and psychic), subject to his limited knowledge of the world and his preferences. This general, and somewhat tautological, characterization of migration nonetheless neglects the dynamic and simultaneous aspects of this process that are not treated here.

Only the individual's labor supply response is considered; conditions in all locations are taken as given in interpreting the individual's selection of an "optimal" move. But as aggregate regional supplies of labor respond over time to differences in economic and social conditions, migration alters these conditions. To disentangle the dynamics of migration some have examined initial period labor market conditions as determining subsequent migration. But a puzzle of modern economic growth, at least in the early stages of development, is the persistence over relatively long periods of migratory patterns and presumably also the conditions that elicited these patterns.* Rankings of regions by economic conditions are, therefore, very similar from decade to decade, and little added insight into causality is obtained by time-ordering aggregate explanatory and dependent variables. Clearly, a dynamic general equilibrium model of the development process is called for to make sense of the determinants and consequences of a time series on migration, in particular, and factor mobility, in general. But models of factor mobility have not yet reached a stage where they are particularly helpful in interpreting data on migration, either at an aggregate or individual level. Therefore, the static, partial equilibrium approach is adopted here, that unfortunately neglects the spacially distinct determinants of the derived demand for labor and the interactions over time among interregional demands and supplies of labor. In this regard, the current framework has the limitations of much of the human capital literature that treats individual behavior as a response to predetermined wage differentials, relative prices, and nonearned income.

*The persistence of regional factor price differentials might be explained by exogenous shifts in regional demands for labor that are persistent through time, or by some unorthodox mechanism by which immigrants enhance factor productivity maintaining differential opportunities for labor rather than closing them. This latter explanation would be consistent with Kuznets' conjecture that migrants may raise the average product of labor, because of their selectivity and adaptability (1964, xxxi+.)

Market Returns to Migration

Since migration requires resources and time to realize a new set of employment and consumption opportunities, it can be treated as an investment opportunity. For ranking and choosing among investments, it is appealing to summarize the associated costs and benefits over time as an internal rate of return or present (discounted) value. Well-known problems of thus ranking physical capital investments are at least as severe when these summary measures are turned to human capital investments, particularly migration.* Whatever summary measure of gain or return is associated with migration, it will unavoidably be a very partial measure of the expected psychic, pecuniary and opportunity costs and benefits, appropriately adjusted for risk.**

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*The gestation period of a human capital investment can be a crucial feature in its attractiveness, and yet plays no distinct role in the above summary measures. The importance of time phasing of inputs and outputs can be attributed to imperfections in the human capital market that largely necessitate self financing, and the inability of investors to diversify commitments to reduce risk, since only one choice of migration destination can be pursued at a time. These features of migration help to explain the prevalence of "stepwise" patterns of migration noted since the Industrial Revolution (Ravenstein, 1885), widespread networks of relatives and extended family that facilitate and mobilize capital for migration in some societies, and the relative infrequency of return migration where substantial costs of relocation and job search are incurred initially by migrants.

**Risk is a dominant element in the migration decision, for which measures are imperfect and possibly misleading. There is not only the risk of pecuniary failure, that would weaken the incentive to any investor,there is also the uncertainty of how fundamental changes in the migrant's mode of life and opportunities will change his values and family attachments. Both risks might restrict a youthful migrant's access to family savings, though I suspect the altruistic obligations that characterize the family assure that the extended family is the primary source of monetized investment funds used in migration. Changes in lifestyle might reasonably be disquieting to the migrant's elders, but the ability to bequeath these locational "benefits" to heirs makes migration unusual as a clear source of intergenerational externalities.

Relative or Absolute Differences in Earnings

Economic logic does not indicate whether migration is likely to respond to the difference or to the ratio of earnings. It is simple to show, however, that this specification choice could depend on whether direct costs or opportunity costs of time are the primary deterrent to migration (DaVanzo, 1972). Neglecting consumption benefits from migration, the present value of migrating from region i to j can be expressed:

$$V_{ij} = \sum_{t=1}^{n} (W_{jt}-W_{it})/(1+r)^{t} - C_{ij} - P_{ij} - T_{ij}W_{i1} \qquad (1)$$

where W_{jt} and W_{it} are the earnings opportunities available to the potential migrant in period t in region j and i, respectively, n is the retirement age minus the migrant's current age, r is a constant discount rate, and C_{ij}, P_{ij} and T_{ij} are the direct, psychic, and time costs, respectively, of migrating from i to j, all of which are assumed to be incurred in the initial period. Time costs are valued, in this example, at the initial period origin wage.

For simplicity assume that regional wages do not vary over time and working ages, t=1,...n; the internal rate of return, r^*, is then the discount rate that equalizes the present value of current costs and annuity benefits.

$$C_{ij} + P_{ij} + T_{ij}W_i = (W_j-W_i)/(r^*(1-r^*)^{-n}) \qquad (2)$$

If we abstract from the finiteness of the working life, and let n approach ∞, then

$$r^* = (W_j-W_i)/(C_{ij}+P_{ij}+T_{ij}W_i). \qquad (3)$$

Assume first that migration costs are only opportunity costs of foregone earnings during the period of relocation and job search, $C_{ij} = P_{ij} = 0$,

$$r^* = (1/T_{ij})\ ((W_j/W_i) - 1), \tag{4}$$

and the migration function might have as its arguments the ratio of wages minus one, and the reciprocal of the time units foregone by migration.

Alternatively, if direct and psychic costs are the only costs of migration, and they were unrelated to origin or destination wages, i.e., $T_{ij}=0$, then the difference in wages might be an argument in the migration decision function with the reciprocal of the direct costs:

$$r^* = (1/(C_{ij}+P_{ij}))(W_j-W_i). \tag{5}$$

Regardless, the internal rate of return is expressed as a product of the arguments representing the cost and benefit components; actual specifications of these terms would, of course, depend on the nature of available data, but the multiplicative form indicates the logarithmic specification of the Z's would have some basis in theory, and my inclination is to view the time costs of migration and job search as the dominant constraint on migration.

Migration costs are usually approximated by the distance from i to j, which leaves much to be desired. And though direct costs, C_{ij} and P_{ij}, may be a well behaved monotonic function of distance, the link to opportunity costs, $T_{ij}W_{i1}$, is unclear. To approximate regional differences in T_{ij} one needs added information on job turnover and an explicit model of how jobs are allocated.

Private Internal Rates of Return to Migration

In contrast with the literature estimating earnings functions (Mincer, 1974), estimates of the parameters of migration functions do not translate, even approximately, into the internal rate of return to migration. But at a more descriptive level I propose that the relative standard deviation (of the logarithms) of wages across regions be considered as an indication of the average magnitude of gains available to migrants in terms of time costs.* Table I presents comparative statistics on male wages and migration across regions of Venezuela.** As in many other settings, migration is observed to increase with educational attainment. In contrast with the tendency for interpersonal relative variation in wages to increase with educational attainment,*** the interregional relative (logarthmic) variation in wages generally diminishes with educational attainment, as seen in Venezuela. Males with some primary schooling in Venezuela report in 1961 an interregional relative variation in wages of .22. With a log normal distribution of regional wages, a representative potential migrant residing in a state with the geometric mean level of wages would find about 16 percent of the alternative regions (i.e., greater than one standard deviation above the mean) offering him a wage at least 22 percent greater than that which he currently receives. For males with some secondary schooling, a similar fraction

*Many additional productive attributes of a labor force might differ across regions and explain interregional earnings differentials. Figures for women in Venezuela (Schultz,1976) and for men in the U.S. (DaVanzo, 1972; Schwartz, 1972) evidence the same pattern.

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*** For the U.S. see Mincer (1974) and Schultz (1971). Evidence from other countries is widely scattered with some exceptions being found within narrow advanced specialities, for example, in Netherlands. But across general educational classes with no less than five years working experience, the tendency for relative variance to increase with education seems common.

Table 1.

Levels and Variation in Male Wage and Lifetime Migration Rates in Venezuela in 1961 among Coterminous States

	Arithmetic		Logarithmic	
	Means	Standard Deviation	Means	Standard Deviation
Monthly Average Wage Rates				
No Schooling	368	145	5.84	.345
Some Primary	558	118	6.30	.220
Some Secondary	1629	201	7.39	.120
Some Higher	6119	530	8.71	.096
Lifetime Average Migration Rates				
No Schooling	1.03	2.26	-1.43	1.74
Some Primary	1.61	3.39	-.808	1.56
Some Secondary	3.03	6.57	-.083	1.48
Some Higher	4.04	8.62	.140	1.57
Nonmigration Rates				
No Schooling	80.4	6.63	4.38	.082
Some Primary	69,7	8.51	4.24	.121
Some Secondary	42.4	12.3	3.71	.276
Some Higher	23.7	15.4	2.97	.590

Source: Schultz, 1976 table 4.

of alternative regions would present him with at least a 12 percent gain, and for those with some higher education, a comparably common gain would be less than 10 percent. If the time costs needed to obtain these destination average earnings streams were equivalent to one year, these percentage gains would also approximate internal market rates of return.*

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*One might expect that regions with very different wages would also tend to be separated by greater distances, or have other countervailing factors responsible for some portion of the wage gap. Another approach to estimating internal rates of return from wage relatives might be to seek to explain this relative variation, and regard the standard deviation of the residual in such a model as a measure of unexplained regional variation that might warrant interregional migration. The residual in such a wage model might then be entered as an argument in a migration decision function.

Figure 1 plots by education these observations on Venezuelan male migration rates, $\bar{m}_{ij}$, against my proxy for the level of returns to migration, σ (ln w_j), namely, the standard deviation in the logarithm of wages across regions. It is not now possible to interpret this direct relationship as evidence of a structural relation that an individual or group would confront, because of the difficulty of identifying the underlying supply and demand schedules. Migration is analogous to schooling in this respect, where it has been noted that marginal returns decline in a cross section with increased levels of education (Becker, 1975, Rosen, 1975). The observed locus, however, involves both individual and group differences in the supplies of funds for investment in schooling or migration, and differences in schooling or migration opportunities (or the demands) for individual and group skills. It seems probable that both the supply of resources and efficiency of and opportunity for migration differ across educational classes, and thus both structural factors have a hand in the reduced form empirical relationship plotted in Figure 1.

Migration research has not yet distinguished between these alternative hypotheses, though their implications for policy are quite different.*

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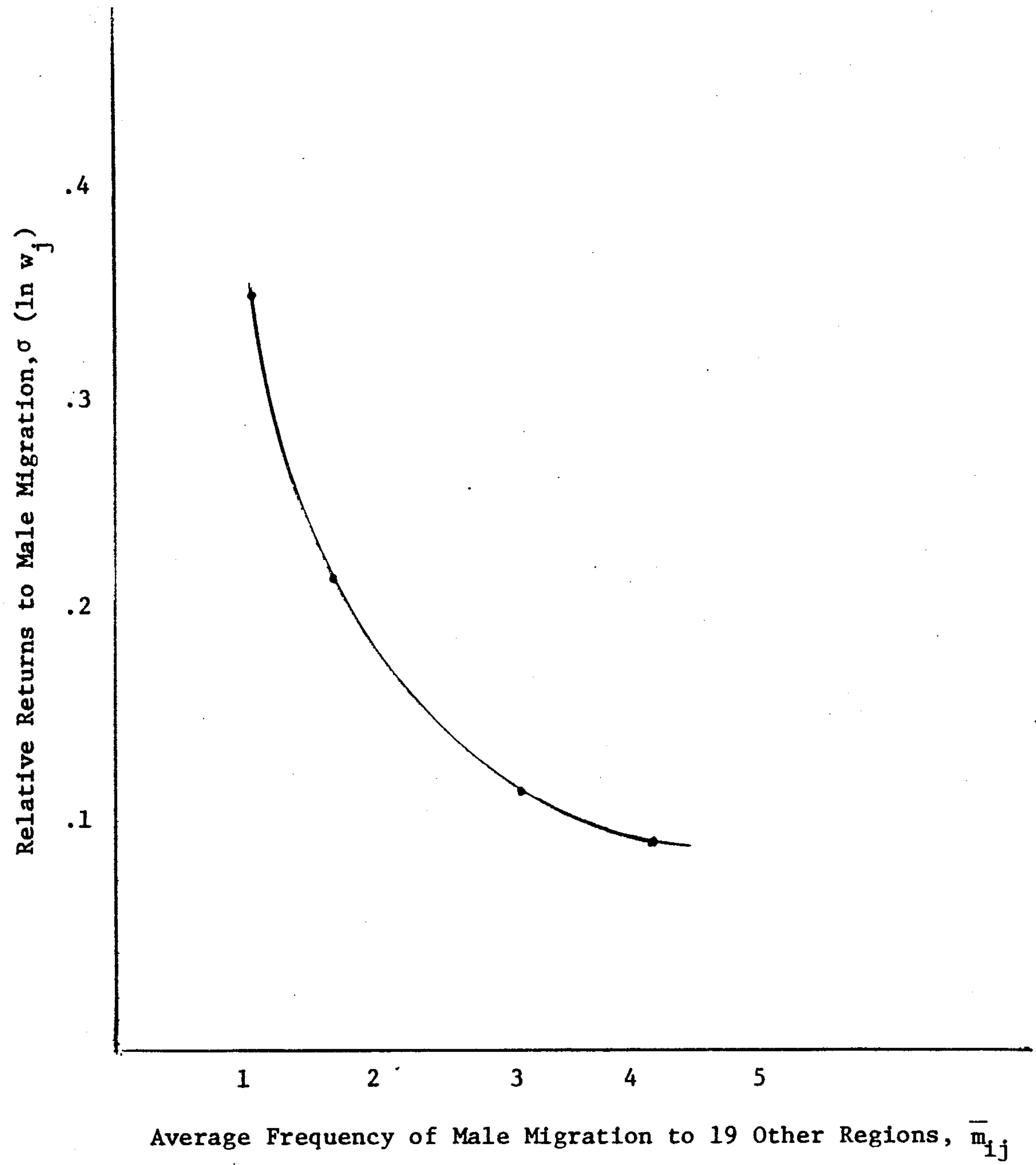
* Consider the extreme of perfect migration capital markets where persons all have access to the same supply schedule of investable funds. Then, the pattern in Figure 1 could be explained by a tendency for the better educated to be more productive in migration (in schooling) in ways not accounted for by my proxy for the returns to migration. For example, lower opportunity costs might be incurred by the better educated if they lose less time in movement and job search, or if neglected nonpecuniary or cultural gains from migration are relatively more important to better educated. No policy rationale would exist, therefore, to narrow the difference in measured returns to migration across education classes. At the other extreme, consider the possibility that the opportunities for migration, or the demand schedule, is identical for all educational groups, and variation in the supply of resources to invest in migration by education group determined the locus of observations in Figure 1. In this instance, differences across educational groups could be viewed as tracing out a common migration investment schedule, in which the increased frequency of migration by the better educated drives down their marginal returns to further migration. According to these assumptions, society might wish to help the less educated avail themselves of high return migration opportunities and thereby reduce interregional and interpersonal income inequality.



FREQUENCY OF AND RETURNS TO MALE MIGRATION IN VENEZUELA

Figure 1

Labor Markets and the Tradeoff between Employment and Wage Rates

Distortions in factor prices contribute to costly misallocation of resources in many countries; there is no reason to think that Venezuela is an exception.* The most frequently stressed factor market distortion is the underpricing of foreign exchange and capital imports that deters a country from exploiting its real international comparative advantage. Conversely, dual labor markets are thought to have counterproductive effects on both the efficient utilization of labor and the equitable distribution of income among persons. Yet in this latter case there is scant empirical evidence on such distortions, and in particular few studies on how these distortions might affect internal migration. This is all the more peculiar since models of development emphasize differences in the functioning of labor factor markets and for this reason prescribe notably different policies with respect to employment expansion and rural-urban development priorities (Todaro, 1969). Thus, in specifying and interpreting estimates of a model of internal migration, an important policy objective is to appraise labor market distortions that might create a divergence between privately and socially optimal levels and patterns of internal migration.

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*For example, from 1960 to 1964 Venezuela maintained a 25% subsidized exchange rate for the importation of machinery and heavy capital equipment, a period of recession and substantial unemployment. Childers, 1974, p. 31.

The salience of urban unemployment led some to doubt that migration was an adequate response to regional variation in employment opportunities. Todaro(1969) and Harris and Todaro(1970) interpret urban unemployment as a market clearing "price" between sectors in a low income country in which urban wages are institutionally maintained in excess of their market equilibrium level.* Rural-urban migration occurs, in the Harris-Todaro (HT) model, until unemployment reaches a level at which appropriately discounted "expected" lifetime earnings are equal in urban and rural sectors. The policy significance of the HT formulation is that it admits to the possibility that added urban employment could <u>reduce</u> social product; the opportunity cost of attracting labor from agriculture might exceed the social product of the new urban job,** given the attendant increment to urban unemployment. This crucial result hinges on the responsiveness of migration and the difference between labor's marginal product in the two sectors. In contrast, traditional models of dualism and development assume that "unlimited supplies of labor" are forthcoming from agriculture at little or no opportunity cost, and consequently, expansion of urban employment is necessarily socially productive.

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*Other explanations have been offered for urban-rural wage differences. Some stress the family organization of production in peasant agriculture which tends to reward the individual on the basis of labor's average product, when the marginal (social) product is less (Sen, 1975, Chap. 6). Others emphasize the need of modern urban firms to reduce turnover and encourage firm specific job training (Stiglitz, 1972). Finally, appeal can be made to differences in tastes for rural and urban livelihoods, and of course, differences in the cost of living.

** The HT model can be generalized in a number of directions, for example, to the case where capital is also mobile between sectors. For a lucid presentation of the framework and its implications for policy see Corden and Findlay (1975).

In the HT model, it is assumed that potential migrants behave as though they maximized their expected earnings, defined as the product of their expected wage rate, and their perceived probability of finding employment, expressed over time and discounted to present values. In determining who gets the available urban jobs, HT assume that all job seekers have an equal chance, and consequently, expected employment in each period is one minus the average unemployment probability. Stiglitz (1972) shows that the same expression holds for the expected urban wage in the absence of urban growth for either the queuing model, in which individuals are hired in the order of urban arrival, or the random selection-poisson model, in which individuals are hired irrespected of their arrival times.

In empirically estimating a migration function in which explanatory variables are expressed in logarithmic form, the expected earnings hypothesis suggests that the coefficients on the logarithm of the wage rate and employment rate are identical.* This is, of course, a severe empirical test of the HT formulation, but relevance of the model or the adequacy of the data may be questioned if the destination employment rate coefficient is not positively and significantly associated with migration (Schultz, 1976).

* This formulation neglects important information on the period of job turnover or the duration of unemployment. Such data would permit more satisfactory testing of alternative models of the labor market in which jobs are allocated over time in a specific manner. Clearly, if employment were randomly allocated <u>each day</u> regardless of arrival in the labor market, the expected income maximization exercise would seem reasonable. If the reallocation of employment occurred at <u>yearly intervals</u>, the risk of being unable to achieve a smooth consumption stream without large wealth holdings provides a rationale for risk averse behavior that might assign greater weight to the employment probability (and duration of unemployment) than to the expected wage rate when employed. In the Venezuelan Census there are, unfortunately, no tabulations on duration of unemployment by the appropriate categories. For a new treatment of labor market turnover and unemployment as they might affect migration, see Fields and Hosek, 1975.

Asymmetry of Origin and Destination Conditions.

Additional considerations suggest that the treatment of employment conditions in origin and destination regions may be asymmetric. Just as the potential migrant may anticipate that he would encounter more than the average unemployment rate at destination, as a new arrival in the city, he may equally well discount origin unemployment, given his established contacts and family ties. Consequently, origin employment coefficients would tend to be distinctly smaller than destination employment coefficients. This appears to be implicit in the HT formulation where rural employment probabilities are ignored or assumed equal to one.

As an illiquid investment in the productivity of the human agent, migration is undoubtedly constrained by imperfection in capital markets. The income or wealth of the potential migrant or his family is likely to augment his supply of investable funds, and contribute to lowering the return he requires to migrate. This investable-funds effect may be captured by origin wages which would offset, to some degree, the origin wage's restraining effect on outmigration. Origin wage variables may be expected to receive, therefore, a somewhat smaller (negative) coefficient in absolute value than will the destination wage (positive).*

Another common characterization of migration involves the selectivity with which migrants are drawn from their origin population. Lee (1966) concludes that when the opportunities of the destination region fuel the migration process, migrants are positively selected, which could imply for our purposes that better

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* This argument is elaborated by DaVanzo (1972) and tested against US interdivisional gross migration flows. Greenwood (1971b, p. 259) found rural origin income effects were even positive on Indian migration to cities. I would expect to find in Venezuela that the capital market constraint would be most frequently binding in the case of the migration of the least educated. Therefore, the ratio of destination to origin wage coefficients should be greatest for this group.

educated migrants should be relatively more responsive to destination variables. Conversely, when deterioration in origin conditions stimulates outmigration, a negative selectivity arises according to Lee, which suggests relatively greater weight should be associated with origin conditions in the migration of less educated groups. To my knowledge this interpretation of the selectivity hypothesis has not been directly documented; testing for the asymmetry of origin and destination labor market effects by education level is a start, though it does not do justice to the subtle dynamic considerations that may be important in Lee's interpretation of historical evidence.

Urban-Rural Sectors

Urban and rural subsectors of regions should be analyzed separately, for the commensurate measurement of employment and real wage rates across the sectors is hardly possible at this time. Employment levels are reportedly high in rural-agricultural regions, and low in urban-industrial regions.* Yet it is commonly assumed that the majority of self-employed workers in agriculture are less fully employed throughout the year than such unrefined census data indicate (Turnham, 1971). On the other hand, the greater frequency of unmonetized payments in kind (See Venezuelan Census) and lower prices of food and housing in the rural sector understate *real* rural wage rates in comparison with urban. Lacking data on Venezuelan migration and wages by urban and rural sectors, the share of a region's population resident in towns of more than 2500 inhabitants is included to control partially for all of the enumerated problems in measurement and omission of variables, as well as the possibility that attributes of urban (or rural) living may in themselves attract (or repel) a potential migrant.

*1961 was a year of recession in Venezuela, with unemployment reaching 13 percent of the labor force or twice the level recorded in 1950 and again in 1969. Yet in agriculture the unemployment rate was about a third of the national average, and in construction and unspecified (largely urban) activities it was twice the average (Childers, 1974, p. 10).

Distances

A traditional proxy for migration costs--psychic, pecuniary and opportunity--is the road distance between the capitals of two states, D_{ij}. Clearly, this distance variable is a surrogate for much more than the transportation costs of migration. For example, the cost of information is likely to increase, and hence the associated risk and uncertainty of migration would increase, with distance. Cultural and language barriers may also be more difficult to overcome as distances mount. One implication of this interpretation of the distance effect is that better educated persons should have a comparative advantage in longer distance moves, being more adept at obtaining and evaluating information on distant job opportunities (Lee, 1970). Cultural differences may also represent less of a hinderance to migration among the better educated, but firm evidence on this score is scarce. If one is willing to view the wage ratio as an indicator of the rate of return to migration net of opportunity costs of time, the distance variable represents only the remaining pecuniary and psychic cost of migration.

School Enrollment Rates and Educational Attainment

Stratification of the migrant population by sex and educational attainment is essential to quantify the diverse effects of schooling on migratory behavior and to explicitly recognize the heterogeneity of the labor force and population. In addition, educational opportunities of a location are often reported by migrants as an important reason for moving, either for their own access to improved schools or for their children's access (Nelson, 1970). Consequently, the primary school enrollment rate for children between the ages of 7 and 14, S, is considered as a measure of the availability in a region's provision of public services.

The Rate of Natural Population Increase

One potentially important determinant of migration that is exogenous from the individual's point of view but may be **somewhat ameanable to social policy is** the difference between regional birth and death rates. Regional differences in this rate of natural increase of the population stimulate migration to the extent that these differences do not correspond with regional employment growth. Population growth has often been greater in rural areas than in urban areas, and these rural areas have also **frequently experienced** slower growth in derived demands for labor. Consequently, both supply and demand shifts have reinforced disequilibrium among labor markets stimulating internal migration. The partial equilibrium framework adopted in this study interpretes employment conditions as motivating individuals to migrate, but does not attempt to determine how these conditions were produced by shifts in regional derived demands for labor and regional differences in natural increase in supplies of labor. Kuznets in his introductory essay to Population Redistribution and Economic Growth, U.S. 1870-1950, concludes that "the effects of population increase are far less important than those of structural changes in the economy's productive system" (1964, xxv). But in understanding contemporary migration in developing countries, regional differences in population increase may no longer be secondary to changes in the structure of production (Schultz, 1969). Though the Venezuelan data are less than ideal to examine this issue, a cursory analysis is attempted.

The age structure of a closed population conveys much information about the rate of natural increase of the population, and even its constituent parts, namely birth and death rates. But in open regional populations, subject to substantial net migration, rates of natural increase and vital rates may not be inferred with great confidence from Census data alone, such as are available from Venezuela. Since most migration occurs among youth, a first approximation for the potential growth in labor supply that is likely to increase migration, other things being equal, is the proportion of the population entering the adolescent stage in the life cycle. Clearly, this proportion cannot be measured ex post in 1961, after migration has taken its toll, but may be better approximated ex ante in the 1950 Census by restricting attention to the child population that has not yet reached the age of most frequent outmigration.

The proxy used here for the natural rate of increase is the percentage of a region's population that is less than age ten as reported in the 1950 Census, denoted by G. As the child death rate fell in Venezuela this variable increased by one-fifth, from 28.1 percent in 1936 to 30.5 in 1950 to 33.8 in 1961. It is assumed in this study that regional differences in vital rates during the 1940's are exogenous to the subsequent pattern of economic development.* The natural rate of increase in a region's labor force is, therefore, expected to deter immigration, other things being equal.

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*To the extent that fertility increased or child mortality decreased in Malthusian fashion as regional employment conditions improved, one might expect to find an endogenous direct relationship between population growth and conditions attracting migrants into a region. Inclusion of this supply shift in a model of migration may reduce the estimated partial effects of employment and wage conditions, to the extent that they are themselves influenced by the regional distribution of population growth, exclusive of migration.

Empirical Specification of an Aggregate Migration Model for Venezuela

The Venezuelan data impose simplifications and empirical compromises, some of which have already been discussed. Both the limitations of the dependent variable and the ambiguous meaning of regional population size require discussion. The dependent variable is an estimate of the probability of migration from region i to region j, P_{ij}, or its aggregate counterpart, the gross migration rate. The denominator in this rate is the number of persons of the specific sex and education group, over seven years of age (in 1961), born in the i^{th} state and enumerated in the universe of 20 coterminous Venezuelan states. The numerator is the number of that specific population who are resident in the j^{th} state in 1961.

This lifetime measure of migration has many undesirable properties (Elizaga, 1965); it can be replaced, however, by an annual (last year) migration rate from the 1961 Venezuelan Census only if the educational breakdown is sacrificed. The prolonged period over which gross migration is measured neglects differential mortality among migrant groups and those that have moved repeatedly, possibly ending up in 1961 outside of the country. Also, birthplace may not be a place of permanent residence, particularly where large municipal hospitals provide maternity services, as in Caracas, for a dispersed population. These measurement errors are probably less serious than the loss of precision in the time dimension. In particular, employment and wage conditions in 1961 may not measure satisfactorily the conditions influencing migration over the prior two decades.*

Associations between the population size of regions and gross migration rates to or from that region are difficult to interpret as evidence on any structural equation because of measurement error and omitted variable bias. In the first case, many "gravity" studies account for past migration in terms of current population size variables. Since migrants are counted in current destination populations and excluded from current origin populations, a positive and negative definitional correlation (bias) is introduced that distorts any time ordered association between migration and population size variables.** This bias is particularly serious when migration is measured over long periods, as here a lifetime, and in settings where migration flows are predominantly in one direction.*** By redefining population size variables ex ante, as the number of persons born in the region, this definitional bias can be removed, but there remains another, more subtle, bias that arises from the persistence of interregional patterns of development and population growth.

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* Some comfort, however, can be drawn from the high correlation between Venezuelan one-year and lifetime migration rates by sex (for all education groups combined). Indeed, when gravity models of migration are estimated using annual and lifetime migration rates, parameter estimates are similar (Levy and Wadycki, 1972a; also for India, Greenwood, 1971a). Substituting Venezuelan employment rates from the 1950 census into the models used here did not change notably the results. Wage information is not available before 1961, however, which precluded a uniform shift of all variables to the earlier period.

** This would appear to be the procedure followed by Beals, et. al. 1967; Greenwood, 1969b, 1971; Levy and Wadycky 1972a, 1974b; and Sahota, 1968, among others.

*** Migration is called "efficient" if the net migration flow from one point to another is large relative to the sum of the gross migration flows occuring in both directions. In low income countries, migration tends to be more efficient (unidirectional), particularly among the less educated. See related discussions by Lee, 1970; Sjaastad, 1962, Schwartz, 1971.

Frequently populous regions are so populated because they contained early centers of commerce, industrialization, and urbanization, and subsequently attracted a net inflow of migrants. When one observes migrants continuing to gravitate toward more populous regions, at least in most low income countries today, this may not be due to the larger number of persons in the destination regions, as implied by the gravity model, but only a persisting reflection of the omitted or imperfectly measured variables that continue to influence migration.* The populous regions once had the prerequisites to amass a large population, and these advantages appear to be eroded slowly, if at all, by the development process. Caution must be exercised, therefore, in interpreting the coefficient on population size variables for it may reflect a "size effect" or the effect of many omitted regionally persistent variables. An improved dynamic approach to migration flows overtime and across regions might disentangle this ambiguity.**

*In several studies the prior stock of migrants has been considered as a determinant of current migration, using single equation estimation techniques. The effect of this variable is rationalized in terms of information flows or the effects of friends and family on migrant destination choice. But in this case, even more clearly than with population size variables, the prior migrant stock is an endogenous variable, and by not treating it with simultaneous equation techniques, the migration equation is seriously biased. Not surprisingly, the prior migrant stock explains very well current migration flows, in both the US and Venezuela. See Greenwood, 1969a and Levy and Wadycki, 1973.

** One way to test this hypothesis concerning the appropriate interpretation of destination population size effects is to pool a time series of cross-sections on interregional migration. The disturbance in the estimated migration equation could then be partitioned into a region specific and random component using the procedure first proposed by Balestra and Nerlove (1966). My expectation is that this more appropriate dynamic estimation approach would "wash out" the effect of both destination and origin population size variables. It would also, in all likelihood, reduce the magnitude of coefficients on other variables that are highly serially correlated overtime in the cross section. See Schultz, 1973.

In the case of Venezuela, the regional configuration of growth in the demand for labor appears to have changed in the 1930's.* The native born population variable, B, (born before 1954) which is used in this study reflects to a great degree the population distribution before the contemporary regional pattern of Venezuelan development took hold of the economy after the Second World War. The case can be made, therefore, that the Venezuelan population size variable does not serve as a proxy for relevant employment conditions, but rather reflects the advantages that accrue to those seeking employment in what were larger (or smaller) sized labor market as of about 1940. It is my expectation that this scale of market effect would be valued most highly by skilled and technically specialized workers.

*For example, Morse (1971 p. 40) cites evidence that the proportion of the Venezuelan population living in the metropolitan Caracas area fell from 6.5% in 1825 to 3.9% in 1881 and rose to only 4.8% by 1920. By 1941 the petroleum and mineral exploitation boom had begun, initiating the current regional configuration of growing and declining employment opportunities. By 1941 9.2% of the Venezuelan population resided in the Caracas metropolitan area, and by 1961 the figure had reached 17.8%.

III. A Statistical Model: The Conditional Logit

My objective is to specify a set of relationships that describe how the mutually exclusive and exhaustive probabilities of locational choice, including the outcome of not migrating, might depend on a set of conditioning variables. One model for such a phenomena is the logistic model as applied in bioassay for a number of years (Mantel, 1966; Cox, 1970) and more recently in economics (McFadden, 1968; Theil, 1969; Nerlove and Press, 1973). In particular the application of Domencich and McFadden (1975) to the study of consumer choice among urban transportation modes is analogous to the problem analyzed here. They provide a rigorous basis for considering individual choice among discrete alternatives within the traditional framework of economic rationality and utility maximization. Differences among individuals in tastes or utility functions are posited in a stochastic form, providing an econometric link between observed discrete choices individuals make and attributes of the alternatives and observable traits of individual decisionmaker.

An individual is confronted with n alternative locations in which to reside, including his origin location (e.g. birthplace) denoted by subscript i. The probability that he resides in location j in a specific time period is assumed to depend on a vector of weighted personal and regional characteristics, Z_{ij}.

$$P_{ij} = \frac{e^{Z_{ij}}}{\sum_{j=1}^{n} e^{Z_{ij}}} ; \qquad (i,j = 1,\ldots,z) \tag{3}$$

where for each region of origin, probabilities sum to one:

$$1 = \sum_{j=1}^{n} P_{ij} \qquad (i=1,\ldots,z) \qquad (4)$$

The ratio of any two probabilities implied by this specification is independent of the characteristics of other (hence, irrelevant) locations. Though this lack of differential substitutability or complementarity between alternatives may be a shortcoming of the polytomous logistic model, this functional specification provides a flexible and symmetric way to treat multiple choice situations and implies a plausible, if not ideal, characterization of the determination of interregional migration.*

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* For example, one suspects that changes in employment opportunities in Baltimore influence the relative numbers of persons from Philadelphia migrating to Washington, D.C. as opposed to New York City. The cross substitution effect of conditions in Baltimore is probably greater on the Washington inflow of migrants than on the inflow to New York City. On the other hand, changes in opportunities in Seattle might leave these specific flows relatively unchanged as assumed in the logit formulation. More generally, how the spacial organization of locations or the geographic spread of information about locations affects patterns of migration is frequently discussed in the literature but has not yet been resolved in a convincing and empirically tractable way. Levy and Wadycki (1974a) recently attempted, in the context of a gravity model of migration, to operationalize Stouffer's (1940) concept of "intervening opportunities" as a determinant of interregional migration. In most low income countries there are relatively few urban centers of growth. Complex heterogeneous interregional migration flows with substantial cross substitution effects may be less of a problem, therefore, for the study of migration in low income countries.

A possible specification of Z_{ij} would be a linear function in natural logarithms* of (i) the pertinent characteristics of the origin and destination regions, X_i and X_j, (ii) the average distance between persons in the two regions, D_{ij}, and (iii) individual traits associated with susceptability to migration, Y_i. Where theoretical guidance on scaling of Y's is limited and the effect of a trait, such as education, is thought to operate in conjunction with the X's and D_{ij}, stratification of the population according to these invarient traits is a promising research strategy (Schultz, 1976). Response parameters across groups defined by such variables as age, sex, and educational attainment may then be tested for equality and reaggregated where parameter differences are negligible.

$$Z_{ij} = \alpha + \sum_{k=1}^{K} \lambda_k \ln X_{ki} + \sum_{k=1}^{K} \gamma_k \ln X_{kj} \qquad (5)$$

$$+ \quad \delta \ln D_{ij} \qquad (i,j = 1,\ldots,z)$$

where α, δ and λ_k, γ_k for k=1,...,K are the 2K + 2 parameters of the migration probability function for each strata of the population. Before exploring restrictions to reduce the number of independent parameters,

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* The logarithmic form of Z_{ij} is preferred for several reasons. First, the expected wage hypothesis later tested posits multiplicative interaction between wage rates and employment rates which is readily translated into parameter restrictions on the logarithmic variables. Second, if opportunity costs are the major costs of migration, the ratio of expected incomes in two regions approximates the return to migration between these regions (DaVanzo, 1972). Third, the empirical literature on migration has generally fit double log linear equations permitting more nearly direct comparisons. Finally, the logarithmic form of Z_{ij} explained more of the variance than other forms I tried, such as a linear form.

structural differences require consideration that might distinguish the processes determining whether a potential migrant leaves his origin location, and if he does migrate, whither he relocates.*

This "uniform" specification of locational choice as a single integrated decision process provides a reasonable framework for considering migration probabilities, P_{ij}, where $i \neq j$, but does not address potential complications that might arise in the case of non-migration, namely P_{ii}, because of discontinuous costs of relocation.

There is also a problem of measurement. If all regions contain the same area and populations, the nonmigrant probabilities might be treated simply as an adding up constraint, implied by equation (4). But if regions differ in size, relatively larger ones would encompass a relatively larger share of all changes in residence within their own boundaries, augmenting the frequency of measured nonmigration. One anticipates, therefore, that origin area or initial origin population would be positively correlated with measured nonmigration, other things being equal.** As noted earlier, the size of an administrative region is also often correlated with unobserved socioeconomic determinants of migration. Thus, both omitted variable and measurement bias is likely to cloud any interpretation attaching to the origin population coefficient.

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* By analogy, studies of labor supply often treat in an integrated statistical framework the determinants of labor force entry and hours of work decisions by means of estimating parameters to a censored linear normal model (Schultz, 1975).

** To control for the origin region size effect on nonmigration, a proxy is defined for the average distance between persons in a region. Assuming regions circular and population uniformly distributed, the square root of the area in square kilometers divided by 2π is introduced into the migration equation to explain nonmigration probabilities, i.e. to explain P_{ii}.

The "two stage" view of migration might be interpreted as violating the logit specification by assuming that some or all response parameters in the process determining nonmigration are distinct, indicated by asterisk:

$$Z_{ii} = \alpha^* + \sum_{k=1}^{K}(\lambda_k^* + \gamma_k^*) \ln X_{ki} \qquad (i = 1,\ldots,z) \qquad (6)$$

whereas the Z_{ij} for $i \neq j$ are still determined according to equation (5).

Another more readily adopted modification to the migration model might be called the "symmetry hypothesis", in which origin and destination conditions are thought to exert equal but opposite effects (elasticities) on the ratio of probabilities of migration, namely $\lambda_k = -\gamma_k$. It follows then that Z_{ij} linearly depends on the ratio of origin to destination conditions:*

$$Z_{ij} = \alpha + \sum_{k=1}^{k}\lambda_k \ell n (X_{kj} / X_{ki}) + \delta \ln D_{ij}, \qquad (i,j = 1,\ldots,z;\ i \neq j) \qquad (7)$$

and if $\lambda_k^* = -\gamma_k^*$, then,

$$Z_{ii} = \alpha^*. \qquad (i = 1,\ldots,z) \qquad (8)$$

Clearly, certain factors may be symmetric and others not; such possibilities can be tested as restrictions on the estimated parameters.

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* If the X's in (5) entered in arithmetic rather than logarithmic form, the "symmetry hypothesis" would imply a differenced form of (7)'

$$Z_{ij} = \alpha + \sum_{k=1}^{k} \lambda_k (X_{rj} - X_{ki}) + \delta \ell n D_{ij} \qquad (i,j = 1,\ldots,z; i = j) \qquad (7)'$$

Estimation

The uniform polytomous logistic model of migration, summarized in equations (3), (4) and (5), can be estimated by maximum likelihood techniques based on individual or grouped data. If the likelihood function converges to a maximum it has been shown to be unique (McFadden 1968; Nerlove and Press, 1973).

Information on migration frequencies are also often tabulated from large surveys or censuses. Cells in which some, but not all, persons at risk of migration move, and thus the expected migration probability for these cells is greater than zero and less than one, the polytomous logit model can be estimated by ordinary least squares regression. In order to impose the n adding up constraints in equation (4) it is convenient to express the migration probabilities as ratios. I propose the convention of treating the nonmigrant probability, P_{ii}, as the normalization factor. Taking logarithms of these probability or odds ratios, one obtains the desired estimation equation that is linear in parameters:

$$\ln (P_{ij}/P_{ii}) = Z_{ij} - Z_{ii} \qquad (i,j = 1,\ldots,z; i \neq j) \qquad (9)$$

which becomes for the "uniform symmetric model":

$$\ln (P_{ij}/P_{ii}) = \sum_{k=1}^{K} \gamma_k \ln (X_{kj}/X_{ki}) + \delta \ln D_{ij} \qquad (10)$$

Aggregate estimates of this form of the logit model provide no direct information on α (no intercept) or β's (distinct origin and destination effects), and implies a "symmetric" ratio treatment of origin and destination conditions.

When the vector of parameters allocating migrants among destinations is allowed to differ from that determining nonmigration, a "two step" hybrid migration model is implicitly being estimated:

$$\ln (P_{ij}/P_{ii}) = (\alpha - \alpha^*) + \sum_{k=1}^{K} (\lambda_k - \lambda_k^* - \gamma_k^*) \ln X_{ki} \tag{11}$$

$$+ \sum_{k=1}^{K} \gamma_k \ln X_{kj} + \delta \ln D_{ij} \qquad (i,j = 1,..,z; i \neq j)$$

A weak test of the hypothesis that migration should be treated as two separate decisions--whether and whither to migrate---can be inferred from estimates of equation (11). If regression coefficients of $\ln X_{ki}$ are of approximately equal absolute values, but opposite sign, to the coefficients of $\ln X_{kj}$, one can impose the restriction of symmetry, replacing the origin and destination variables with their ratio as in equation (10). However, F tests of coefficients equality (Fisher, 1971), would not actually test whether both $\lambda_k = \lambda_k^*$ and $\gamma_k = \gamma_k^*$, but only test whether $(\lambda_k - \lambda_k^* - \gamma_k^*) = \gamma_k$. Also, the standard t test on the intercept, $\alpha - \alpha^*$ may be informative. A negative intercept suggests a tendency for nonmigration to occur more frequently than predicted by the uniform model and vice versa. The existence of such an "enertia" (negative) or "wanderlust" effect (positive) is admittedly a highly indirect means for distinguishing between the "uniform" and "two-stage" migration formulations, but it is beyond the scope of this paper to consider more formal approaches to this topic.

This brings us full circle to the conventional double logarithmic estimation equation (1) for gross migration rates that may be interpreted in terms of the gravity model (2). As the interval of time diminishes over which migration is measured, P_{ii} approaches unity, and $\ln(P_{ij}/P_{ii})$ approaches $\ln P_{ij}$.

Though the gravity model does not make explicit use of the information contained in the relative frequency of nonmigration, P_{ii}, as does the logit model, the similarity in estimation equations (1) and (11) suggests to me that the double logarithmic equation will often imply reasonable residual estimates for P_{ii}, just as the logit must.* With regard to the empirical specification of X's, the gravity model includes as parameters elasticities with respect to origin and destination populations. The ambiguous interpretation of origin population coefficients was discussed above, and it is also difficult to derive from behavioral assumptions why N_j should be proportional to M_{ij} (Niedercorn and Bechdolt, 1969), though it may be argued that skilled migrants would favor larger labor markets in which heterogeneous demands for specialized labor reduces employment risks.**

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* The coefficient of determination (R^2) is not immediately useful for comparing the fit of the logit and "gravity" models of migration, since their dependent variables differ. The logit estimates of equation (5) can be readily converted into predicted values for all P_{ij} and these compared with the n^2 observed values. Similarly, estimates of the gravity model obtained from equation (1), and the implied estimates for P_{ii}, $i=1,\ldots n$, (though not necessarily positive values) can be compared with observed gross migration rates.

** Greenwood (1971b) even went so far as to assume that the elasticity of migration with respect to origin and destination population size was minus and plus one, respectively, and imposed this normalization on the migration flow to obtain his dependent variable. Vanderkamp (1971) divided the migration flow by the sum of origin and destination populations. Most applications, however, estimate independently the coefficient on destination population size and generally obtain positive elasticities of less than one (see for example, Beals, et. al., 1967; Levy and Wadycki, 1972ab, 1974.)

It should be clear, nevertheless, that given the various enumerated measurement problems with regional population units, and the omitted variable bias inherent in the dynamic aspect of migration, that a fool-proof scheme will not be devised easily to test the appropriateness of the "gravity" normalization of migration, the constant returns to scale restriction,* or even the aggregate convention of treating gross migration rates as a consistent estimator of underlying individual migration probabilities.

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*There are reasons to anticipate that as the size of a region grows relative to others, a point will be reached where the rate of outmigration from the relatively large region will diminish, and given the same incentives to migrate, the rate of outmigration from a relatively small region will increase. But if all regional populations change by the same proportion, gross migration rates might be scale neutral. In this regard, the Cobb-Douglas specification for the gross migration function might be considered where the parameter restriction in equation (1) $\beta_1 = 1-\beta_2$ is tested:

$$M_{ij} = N_i^{1+\beta_1} N_j^{\beta_2} f(Z's),$$

for which the gross migration rate is

$$m_{ij} = N_i^{1-\beta_2} N_j^{\beta_2} f(Z's).$$

IV. Aggregate Evidence of Migration in Venezuela

Patterns of Net and Gross Migration in Venezuela

According to estimates by Chen (1968), the regional pattern of net migration in Venezuela changed relatively little from 1936-40 to 1961. Net outflows persisted from seven states (See Figure 2) extending from the Andes to the central coast (i.e. Táchira, Merida, Trujillo, Lara, Yaracuy, Cojedes, and Falcón) and from two states on the North East coast (Sucre and Neuva Esparta). Conversely, the major urban-industrial centers (Federal District-Caracas; Zuila-Maricaibo; Portugese-Guanare; and Anzoátegui-Barcelona), attracted steadily a net inflow of migrants in about equal numbers from the other states of Venezuela and from abroad, with the metropolitan growth of Caracas spilling over into neighboring states after 1941 (Miranda, Aragua, and Carabobo). The remaining states lost and gained population in various periods, but some sustained substantial net outmigration recently (Monagas and Apure).

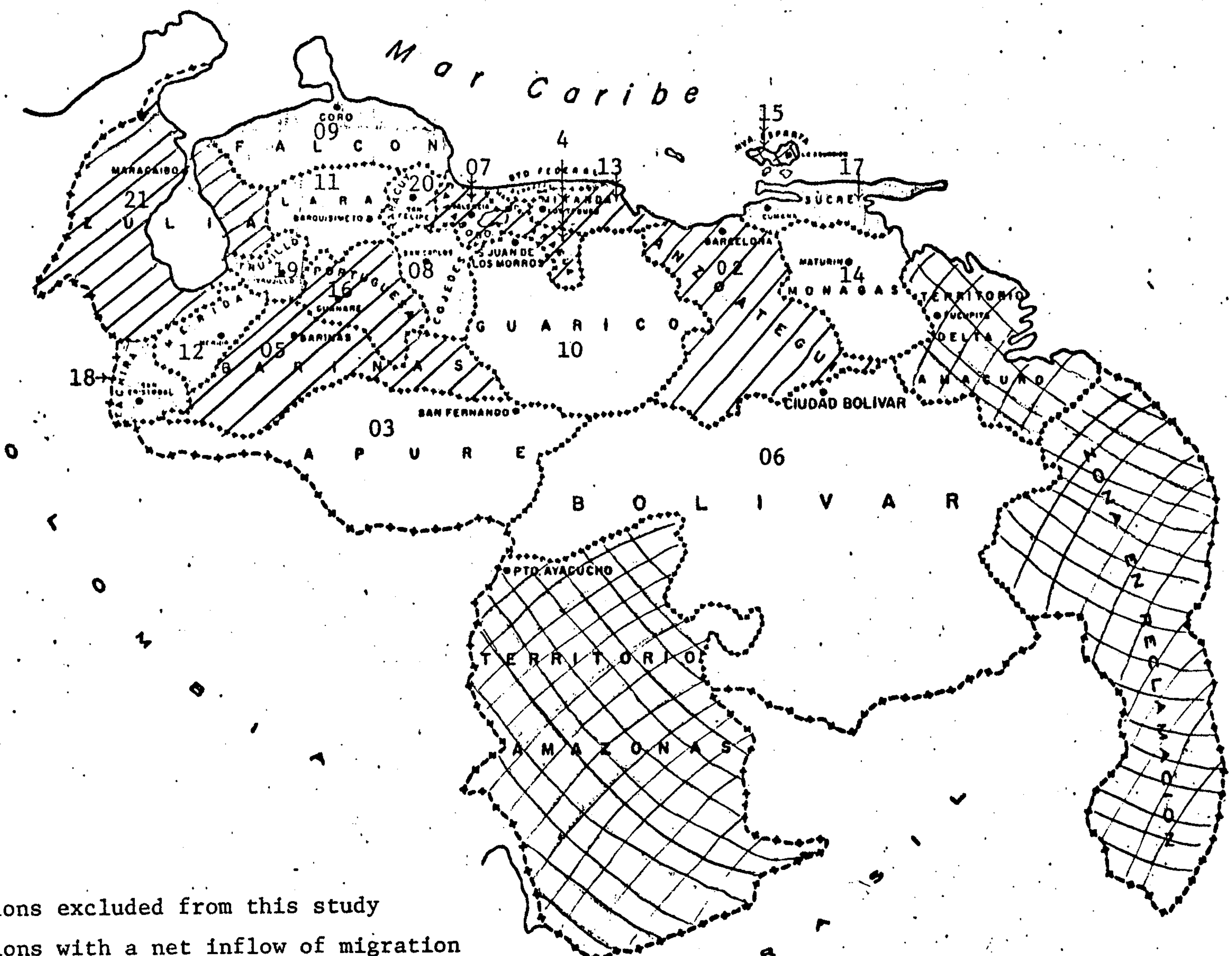
In estimating a closed model of gross migration, that is, one in which the gross rates for each origin population sum to one, one must exclude persons born in the country who left by the time of enumeration and immigrants born outside of the country who currently resided within it. Although international immigration has been an important source of urban labor force growth in Venezuela in the 1940's and 1950's, there is no way to include these flows unless the origin populations from which they emigrated are also included, which is impractical.*

Gross migration rates for 20 coterminous states of Venezuela

*To this, omitted "rest of the world" sector, for data reasons I have added the sparcely populated territories of the Amazon, Amacuro Delta, and Federal Dependencies, as well as the island of Neuva Esparta. These omitted areas of Venezuela contained less than two percent of the total population enumerated in the Venezuelan 1961 Census.

Figure 2

POLITICAL DIVISIONS OF VENEZUELA



- regions excluded from this study
- regions with a net inflow of migration from other states and from abroad
- regions with a net outflow of migration from other states and, from abroad

Source: Chen (1968).

are reported in Table 1, by sex and by four educational attainment groups. These rates are the proportions of the population born in the state over the age of 7 that are enumerated in another state. The stratification of migration by educational attainment appears promising, for within each sex group and in every state, gross migration rates increase with educational attainment. Differences by sex within an educational attainment group, however, do not follow an obvious pattern. In general, those states that have experienced net immigration have somewhat lower gross (out) migration rates, as one might expect, but the education specific rates are more complex; opportunities for different educational-skill groups do not appear to have grown at a uniform rate in all regions. Table 2 reports the difference between offsetting gross flows which expresses the net inflow (positive) or outflow (negative) of migrants as a percent of the population born in the region with a specific education. These net migration "rates" are not to be interpreted as a migration "risk" or probability for any specific individual or group.

Interregional Variation in Wages and Employment

It has been frequently noted in studies of migration that the better educated appear to respond to interregional differences in income more readily than do less educated (Lee, 1970; Schwartz, 1971; DaVanzo, 1972). It was observed earlier that the relative (logarthmic) variation in wages and incomes across regional labor markets tends to diminish with increasing educational attainment.

Table 1

Percentage of Persons 7 Years and Older, Not Residing in 1961 in Their State of Birth, by Sex and Educational Attainment: The Sum of Gross Out-Mirgation

		Men				Women			
Code	State	No Schooling	Some Primary	Some Secondary	Some Higher	No Schooling	Some Primary	Some Secondary	Some Higher
01	Federal District	17.26	20.44	28.85	42.25	16.96	19.81	27.45	30.80
02	Anzoátegui	15.21	24.97	59.50	83.91	16.00	25.49	47.48	81.58
03	Apure	17.86	32.31	77.85	95.17	20.43	35.18	73.23	97.80
04	Aragua	22.18	28.80	52.67	83.22	25.75	29.45	45.88	72.00
05	Barinas	13.21	23.03	67.81	86.27	16.14	26.15	62.27	92.54
06	Bolívar	13.34	22.81	58.95	81.78	16.52	27.55	56.28	80.97
07	Carabobo	18.50	24.49	43.42	64.27	21.52	26.19	37.09	67.24
08	Cojedes	24.50	39.28	62.23	87.59	30.20	40.54	56.56	100.00
09	Falcón	29.34	37.20	56.69	79.35	27.80	30.46	44.11	59.59
10	Guárico	17.06	30.00	60.21	82.70	20.96	32.39	47.76	93.55
11	Lara	26.70	30.51	46.26	70.76	24.76	26.19	34.25	71.10
12	Mérida	19.57	36.04	65.57	74.51	23.20	37.73	51.14	66.80
13	Miranda	28.60	41.81	56.28	49.21	34.21	42.90	47.91	58.30
14	Monagas	18.72	34.63	63.67	85.90	22.03	34.97	50.21	90.70
15	Nueva Esparta*								
16	Portuguesa	10.92	22.75	57.26	83.33	13.86	25.64	50.00	84.06
17	Sucre	23.50	41.15	69.47	86.57	25.00	39.36	56.03	89.46
18	Táchira	16.82	29.95	62.35	84.46	22.71	32.19	48.68	89.95
19	Trujillo	28.77	40.73	67.62	85.50	29.77	39.31	54.47	91.63
20	Yaracuy	29.41	43.14	72.11	91.09	34.46	44.06	50.07	87.23
21	Zulia	5.25	9.78	29.92	42.13	4.86	9.32	27.29	48.07

*Nueva Esparta, a small island of 89,492 persons, is excluded from this study for a variety of reasons, as are the Amazon, Amacuro Delta and Federal Dependent Territories.

Table 2

The Net Addition (or Reduction-) in Persons 7 Years and Older Due to Migration, By Sex and Educational Attainment, Expressed as a Percentage of Those Born in The State

Code	State	Men				Women			
		No Schooling	Some Primary	Some Secondary	Some Higher	No Schooling	Some Primary	Some Secondary	Some Higher
01	Federal District	64.04	65.17	54.82	71.17	68.33	41.30	69.75	119.31
02	Anzoategui	10.64	-20.18	-53.78	4.79	10.59	-11.22	-64.91	5.95
03	Apure	-6.47	-65.07	-80.42	-6.84	-27.26	-66.58	-93.40	-11.20
04	Aragua	20.83	37.57	30.24	16.36	17.82	19.19	-15.00	16.45
05	Barinas	18.21	-35.52	-66.55	44.00	0.83	-42.10	-73.14	32.02
06	Bolivar	4.31	-37.47	-58.27	10.73	-7.10	-41.07	-72.12	-0.64
07	Carabobo	14.14	1.23	-17.36	14.89	4.73	-0.82	-26.44	12.97
08	Cojedes	-19.83	-36.01	-68.28	-4.51	-26.08	-36.21	-89.66	-14.92
09	Falcon	-32.33	-42.35	-59.58	-26.73	-25.74	-35.20	-50.61	-24.97
10	Guarico	-6.34	-32.28	-65.67	-2.75	-18.14	-31.40	-82.95	-8.74
11	Lara	-19.52	-18.18	-46.90	-20.79	-13.66	-8.56	-53.44	-18.56
12	Merida	-27.16	-32.05	4.05	-11.25	-29.30	-28.65	41.7	-16.63
13	Miranda	-5.77	120.24	653.48	-11.80	7.24	186.68	423.76	-12.31
14	Monagas	-9.11	-42.03	-64.47	13.38	-13.58	-30.87	-84.89	9.14
15	Nueva Esparta*								
16	Portuguesa	47.08	0.62	-47.03	65.31	17.39	-7.63	-72.47	46.48
17	Sucre	-37.73	-60.71	-77.86	-22.07	-36.60	-48.87	-85.41	23.51
18	Tachira	-24.72	-50.11	-76.05	-12.19	-27.32	-39.36	-41.89	-19.33
19	Trujillo	-35.26	-55.45	-76.61	-25.69	-33.95	-41.86	-84.14	-27.22
20	Yaracuy	-26.48	-50.60	-81.01	-12.30	-31.01	-39.13	-80.14	-20.13
21	Zulia	22.06	-6.02	-13.27	26.39	18.35	-6.10	-26.24	36.33

*Nueva Esparta, a small island of 89,492 persons, is excluded from this study for a variety of reasons, as are the Amazon, Amacuro Delta and Federal Dependent Territories.

If one associates interregional dualism with relative variation in wages for comparably educated men, the extent of dualism in the Venezuelan labor market is much smaller for the better educated. For example, in the Federal District of metropolitan Caracas the average wage of the unschooled male is 833 Bolivars per month, compared with 232 B. for the same education group in the rural state of Táchira. In contrast, the monthly wage of men with some higher education is 5851 B. per month in the Federal District of Caracas and 4263 B. in Táchira (See Data Appendix A-1). Although the absolute gain for the higher educated would be three times that for the unschooled, the opportunity costs of similar amounts of time lost from work in the origin state would be twenty times greater for the better educated than for the unschooled. Thus, dualism and the disequilibrium among regional labor markets exists in Venezuela primarily for the less educated, and is presumably reduced by the more frequent migration of the better educated.

Estimates of Migration Functions

The maximum number of cross migration flows among all 20 coterminous states of Venezuela is 380 (20 times 19), but in all but the primary education group there are pairs of states for which the gross migration flow in one direction is zero, and these observations are initially omitted.* The variables are defined in Table 3. Estimates of the conditional logit model of migration, are calculated by ordinary least squares for four educational attainment groups of males, and reported in Table 4a-d. The first regression includes the explanatory variables reported by Levy and Wadycki (1974b) in their estimates of a gravity model, with the following modifications: (a) the current population is appropriately replaced by the population born in the respective state; (b) the unemployment rate is replaced by the employment proportion; (c) the average wage by education group is estimated using an extrapolation formula implied by the Pareto distribution, rather than fixing the average wage for the open-ended interval at its lower limit (see Data Appendix); and (d) origin area and population growth are included.

The second regression imposes the restriction that migrants respond to the "expected" wage, strictly according to the HT framework, defined as the product of the labor force employment probability and the education specific monthly wage rate (indicated by an asterisk). The third and fourth regressions are analogous to (1) and (2), but impose the symmetry hypothesis that the coefficients on origin and destination attributes are of equal absolute value but of opposite signs (i.e. the ratio of j to i is included); symmetry is imposed for the wage rate, the employment rate, the expected wage, and the urban share, but not for school enrollment, population size, or population growth where the rationale is weaker and the evidence conflicts with the restrictions.

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*The logarithm of the zero gross migration rate is undefined. It is proposed in the statistical literature that in such cases a low value be entered to retain information in zero cells, and this procedure was explored by assuming in all cases where no migration is observed that one migrant moved. With the full sample size of 380 for all education groups the estimates did not change appreciably. Maximum likelihood methods are later used to retain this information without such ad hoc procedures. See Cox, 1970.

Table 3.

The Definitions of Variables in Migration Model*

M_{ij}	The number of persons age seven and older in 1961 born in region i and residing in 1961 in region j, by sex and educational attainment.
N_i	The number of persons age seven and older in 1961 born in region i and residing in 1961 in all j regions, j=1,...,n; by sex and educational attainment. $\sum_{j=1} M_{ij} = N_i$.
M_{ij}	The gross migration rate by sex and educational attainment; M_{ij}/N_i.
D_{ij}	The distance in kilometers from the capital of region i to the capital of region j, following major roads.
A_i	The area of origin region or approximate average distance between persons in kilometers in region i; $\sqrt{\text{Area in square kilometers}/2\pi}$.
S_i	The percentage of children between the ages of 7 and 14 enrolled in school in region i in 1961.
U_i	The percentage of the population residing in urban areas in region i in 1961.
E_i	The proportion of the civilian labor force employed in region i, by sex, but for all educational attainment groups together in 1961.
W_i	The estimated monthly wage rate for wage and salary workers age ten and over in region i by sex and educational attainment in Bolivars in the month preceding the 1961 Census. See Data appendix for Pareto extrapolation procedure used for estimating wage rate for open ended intervals.
EW_i	The expected earnings per month, or the employed proportion times the average wage rate; W_i*E_i .
G_i	The percentage of the 1950 Census population less than age ten in region i.

*Data sources are reported and wage rates derived in an appendix available from the author on request.

Table 4a

ESTIMATES OF THE POLYTOMOUS LOGISTIC MODEL OF LIFETIME MIGRATION FOR MEN, WITH NO EDUCATION

	(1)	(2)	(3)	(4)
Number of Persons Born in i, (B_i)	.0933	.0815	.0307	-.00693
	(.67)	(.62)	(.24)	(.06)
Number of Persons Born in j, (B_j)	.220	.252	.140	.165
	(1.58)	(1.89)	(1.11)	(1.35)
Distance from i to j, (D_{ij})	-1.59	-1.59	-1.61	-1.60
	(16.9)	(17.0)	(17.7)	(17.6)
Area of origin (A_i)	.539	.512	.687	.632
	(3.00)	(3.31)	(4.58)	(4.47)
Schooling in i, (S_i)	4.28	4.10	5.43	5.09
	(3.31)	(3.66)	(6.02)	(6.02)
Schooling in j, (S_j)	-2.93	-2.69	-2.39	-2.07
	(2.79)	(2.68)	(2.73)	(2.51)
Urban Share in i, (U_i)	-.891	-.913	Ratio j/i .750	.822
	(2.60)	(2.78)	(3.22)	(3.68)
Urban Share in j, (U_j)	.644	.721		
	(1.89)	(2.20)		
Employment in i, (E_i)	-.224	-	Ratio j/i -.327	-
	(.08)		(.19)	
Employment in j, (E_j)	.0919	-		
	(.04)			
Wage rate in i, (W_i); or if * = Expected wage, (EW_i)	-.857	-.880*	Ratio j/i 1.39	1.51*
	(1.96)	(2.08)	(4.71)	(5.50)
Wage rate in j, (W_j); or if * = Expected wage, (EW_j)	1.83	1.98*		
	(4.33)	(5.29)		
Natural increase of Population in i; G_i	-2.56	-2.53	-3.98	-4.09
	(1.67)	(1.66)	(3.07)	(3.17)
Natural Increase of Population in j; G_j	-2.02	-1.61	-2.64	-2.31
	(1.65)	(1.46)	(2.43)	(2.21)
Constant ($\alpha-\alpha^*$)	3.48	.777	9.15	8.84
	(.34)	(.10)	(1.43)	(1.38)
R^2	.635	.634	.631	.630
Sum of Squared Residuals	421.90	422.74	425.98	427.33

Sample Size is 379

Table 4b

ESTIMATES OF THE POLYTOMOUS LOGISTIC MODEL OF LIFETIME MIGRATION FOR MEN, WITH SOME PRIMARY EDUCATION

	(1)	(2)	(3)	(4)
Number of Persons Born in i, (B_i)	.0499 (.45)	.0339 (.34)	.0543 (.54)	.0148 (.16)
Number of Persons Born in j, (B_j)	.423 (3.88)	.463 (4.57)	.414 (4.20)	.445 (4.67)
Distance from i to j, (D_{ij})	-1.27 (16.9)	-1.26 (16.9)	-1.30 (17.9)	-1.29 (17.9)
Area of origin (A_i)	.300 (2.25)	.272 (2.40)	.376 (3.31)	.326 (3.08)
Schooling in i, (S_i)	2.13 (2.10)	1.96 (2.16)	2.97 (3.89)	2.70 (3.69)
Schooling in j, (S_j)	-1.67 (1.90)	-1.45 (1.71)	-1.11 (1.47)	-.854 (1.18)
Urban Share in i, (U_i)	-.940 (3.68)	-.961 (3.95)	Ratio j/i 1.05 (6.05)	1.12 (6.83)
Urban Share in j, (U_j)	1.17 (4.57)	1.26 (5.29)		
Employment in i, (E_i)	-.464 (.20)	-	Ratio j/i -.359 (.24)	-
Employment in j, (E_j)	-.516 (.25)	-		
Wage rate in i, (W_i); or if * = Expected wage, (EW_i)	-1.17 (2.92)	-1.22* (3.46)	Ratio j/i 1.26 (4.62)	1.43 * (6.13)
Wage rate in j, (W_j); or if * = Expected wage, (EW_j)	1.36 (3.41)	1.59* (4.82)		
Natural increase of Population in i; G_i	.448 (.47)	.480 (.51)	.0299 (.03)	.0390 (.04)
Natural Increase of Population in j; G_j	-4.10 (5.06)	-3.87 (4.98)	-4.20 (5.42)	-4.03 (5.29)
Constant ($\alpha-\alpha^*$)	4.73 (.62)	2.21 (.39)	2.75 (.54)	2.43 (.48)
R^2	.723	.722	.721	.720
Sum of Squared Residuals	261.24	262.08	262.44	263.47

Sample Size is 380

Table 4e

ESTIMATES OF THE POLYTOMOUS LOGISTIC MODEL OF LIFETIME MIGRATION FOR MEN, WITH SOME SECONDARY EDUCATION

	(1)	(2)	(3)	(4)
Number of Persons Born in i, (B_i)	-.0189 (.26)	-.0310 (.42)	-.0516 (.71)	-.0468 (.65)
Number of Persons Born in j, (B_j)	.878 (12.0)	.858 (11.7)	.842 (11.7)	.838 (11.7)
Distance from i to j, (D_{ij})	-.913 (13.9)	-.917 (13.9)	-.891 (14.0)	-.894 (14.1)
Area of origin (A_i)	.361 (3.31)	.247 (2.70)	.188 (1.97)	.218 (2.49)
Schooling in i, (S_i)	2.55 (2.75)	1.70 (2.05)	.878 (1.22)	1.06 (1.55)
Schooling in j, (S_j)	-.862 (1.10)	-1.40 (1.83)	-1.77 (2.54)	-1.93 (2.88)
Urban Share in i, (U_i)	-1.26 (5.81)	-1.46 (7.35)	Ratio j/i 1.25 (8.41)	1.19 (9.08)
Urban Share in j, (U_j)	1.22 (5.64)	.946 (4.99)		
Employment in i, (E_i)	1.48 (.85)	-	Ratio j/i 2.99 (2.68)	-
Employment in j, (E_j)	6.06 (4.01)	-		
Wage rate in i, (W_i); or if * = Expected wage, (EW_i)	-1.79 (3.74)	-1.94* (4.06)	Ratio j/i 2.20 (6.64)	2.15* (6.62)
Wage rate in j, (W_j); or if * = Expected wage, (EW_j)	2.67 (5.62)	2.39* (5.12)		
Natural increase of Population in i; G_i	2.26 (2.80)	2.49 (3.08)	2.29 (3.86)	2.94 (3.82)
Natural Increase of Population in j; G_j	-4.96 (-7.07)	-5.32 (7.63)	-4.93 (7.27)	-5.00 (7.43)
Constant (α-α*)	-10.09 (1.56)	1.17 (.22)	4.78 (1.09)	4.95 (1.13)
R^2	.797	.791	.789	.789
Sum of Squared Residuals	185.64	190.86	192.40	192.73

Sample Size is 378

Table 4d.

ESTIMATES OF THE POLYTOMOUS LOGISTIC MODEL OF LIFETIME MIGRATION FOR MEN, WITH SOME HIGHER EDUCATION

	(1)	(2)	(3)	(4)
Number of Persons Born in i, (B_i)	-.386 (3.73)	-.392 (3.82)	-.422 (4.16)	-.411 (4.08)
Number of Persons Born in j, (B_j)	.685 (6.33)	.665 (6.18)	.650 (6.15)	.642 (6.09)
Distance from i to j, (D_{ij})	-.656 (6.73)	-.664 (6.83)	-.615 (6.56)	-.621 (6.64)
Area of origin (A_i)	.282 (1.73)	.226 (1.69)	.0957 (.68)	.153 (1.19)
Schooling in i, (S_i)	.704 (.53)	.267 (.23)	-1.70 (1.71)	-1.32 (1.44)
Schooling in j, (S_j)	4.03 (3.66)	3.45 (3.33)	2.52 (2.63)	2.18 (2.44)
Urban Share in i, (U_i)	-1.36 (3.99)	-1.48 (4.83)	Ratio j/i 1.08 (4.67)	.965 (4.80)
Urban Share in j, (U_j)	.795 (2.41)	.524 (1.87)		
Employment in i, (E_i)	-.230 (.09)	-	Ratio j/i 3.88 (2.34)	-
Employment in j, (E_j)	5.97 (2.70)	-		
Wage rate in i, (W_i); or if * = Expected wage, (EW_i)	-1.92 (2.66)	-2.03* (2.86)	Ratio j/i 2.41 (4.77)	2.31 * (4.66)
Wage rate in j, (W_j); or if * = Expected wage, (EW_j)	2.93 (4.00)	2.69* (3.75)		
Natural increase of Population in i; G_i	3.06 (2.49)	3.19 (2.61)	4.11 (3.49)	4.06 (3.45)
Natural Increase of Population in j; G_j	-6.11 (5.67)	-6.51 (6.22)	-5.81 (5.57)	-5.97 (5.81)
Constant $(\alpha-\alpha^*)$	-17.5 (1.42)	-7.91 (.73)	1.32 (.21)	1.70 (.27)
R^2	.681	.678	.674	.673
Sum of Squared Residuals	370.08	373.19	378.74	379.82

Sample Size is 358

The only unexpected sign is that on the area of the origin state, A_i, which was introduced to control for the fact that a greater share of internal migration, defined as all changes of residence, occur within a geographic unit the larger the area of that unit (See Stouffer, 1940 for a compatible interpretation). Thus, the anticipated effect of origin area on a region's outmigration rate is negative. The estimates indicate, on the contrary, that the size of states in Venezuela is apparently not independent of socioeconomic factors related to migration; the largest states are the least dense, frontier areas from which outmigration has been relatively rapid among all educational groups.

Replacing the current population by the native born population tends to decrease the size of the regression coefficients on the destination population size variable, as anticipated (not reported). The effect is most substantial for the least educated, for whom migration is most unidirectional. However, even with the native born population variable, the widely observed "gravity" effect is robust, though its interpretation, as stressed earlier, is clouded and may be a source of parameter bias. The elasticity of migration with respect to destination native born population size increases from .2 for the unschooled to .9 for those with some secondary and .7 for those with some higher education. Apparently, the more educated are more attracted to the more populated regions.

The effect of origin native born population size is implicitly removed in the gross migration formulation, for it is assumed that all individuals have the same probability of migration, regardless of the populousness of their birthplace. But this convenient probabilistic approach does not fit the behavior of the higher educated group, who have a pronounced tendency to migrate less from the more populous states than their "expected" rate, given their other characteristics. This may well reflect the added attraction of larger labor markets for the highly educated, or the unspecified tendency of the better educated to reside in a few concentrated areas, such as Metropolitan Caracas and Maricaibo.

Distance has the large deterrent effect on migration found in other investigations of this genre, with the elasticity falling by more than one-half from the least to the most educated groups, confirming the tendency for the educated to migrate long distances more readily. (See Lee, 1970; DaVanzo, 1972).

School enrollment presents a puzzle. All educated groups tend to migrate out of regions with higher primary school enrollment rates, which could be viewed as an information effect of community education level or simply a reflection of diversity in educational accomplishments of persons within the broad educational attainment groups analyzed here. This is further evidence that the more educated migrate more readily, other things being roughly equal. But more curious is the tendency of the less educated to avoid destinations with higher enrollment rates.* Whatever the reason, only the higher educated group is prone to move toward regions with higher school enrollment rates, other things equal.

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*Nor does the hypothesis that the less educated are at a disadvantage in the job market and thus fare worse in regions with more widespread schooling get much support from these data; the simple correlation between enrollment rates and wages is highest for the wage rates of unschooled and primary schooled males, .69 and .52, respectively, whereas it falls to .29 and -.42 for the wage rates of males with some secondary and higher education.

It is an unfortunate aspect of the Venezuelan data that one cannot disaggregate urban and rural populations while maintaining the detail on educational attainment, sex and wage rates. One procedure to control for differences in real opportunity costs and gains of migration between rural and urban labor markets as characterized by measured wage rates and employment rates is to include a variable for the proportion of the population residing in urban areas. The coefficient on urban share is several times its standard error in all education groups, both at destination (positive) and origin (negative), and of approximately the same absolute magnitude. The elasticity estimates are greater for primary educated than for the unschooled, and greater for the secondary and higher educated than for those with some primary.

The measure of population growth considered here is strongly associated with migration. It fosters outmigration at origin for secondary and higher educated, and among all education groups it repels migration into a destination labor market. The only pecularity in these findings is the unanticipated effect of origin population growth upon the migration of the unschooled. There is a tendency for this least educated group to be unwilling or unable to migrate out of regions that exhibit more rapid rates of population increase. With individual data, one might explore whether this result reflects the effect on mobility of the number of siblings, educational attainment, timing of marriage and own family size (See Caldwell, 1968; Hay, 1974).

The employment rate at origin is never statistically significantly related to migration at conventional levels (two tailed 5 percent). Destination employment conditions, that are stressed in the rural-urban context of the HT model, are positive and statistically convincing only for the secondary and higher educated groups. These data do not reveal that unschooled and primary educated migrants respond to the destination level of employment in addition to their notable response to destination wage levels.*

Tests of Restrictions on the Migration Function

Origin and destination wage coefficients are of the anticipated sign, and the destination elasticities exceed in absolute value the origin elasticities. With one exception, the wage elasticities also increase modestly with education level. In three out of four education groups the restriction that the coefficient on the employment rate and the wage rate are equal cannot be rejected by the F test (Table 5 row 1). Though employment rate coefficients are not significantly different from zero for the two least educated groups, the "expected" wage is more strongly related to migration than the wage rate itself, by which I mean the t ratios increase for the compound variable.

For the secondary and higher educated, for whom the employment rate coefficient is larger than that on the wage rate, migration behavior is broadly consistent with the HT model of factor market distortions. These better educated groups, moreover, appear to be strongly risk averse. The evidence that employment conditions govern migratory behavior of the less educated, however, is not persuasive. This might be because these less educated migrants lack accurate information about employment conditions at destination, or more likely, because the traditional urban sector employment opportunities for the less educated reduce substantially the opportunity cost of searching for a well paid job in the modern sector. As a consequence, the less educated report

*This evidence is qualified by problems of adequately measuring at origin real wages and actual employment opportunities for the majority of poor agricultural workers. It is often conjectured that the rural family agricultural production unit may have difficulty efficiently utilizing labor if its members are rewarded according to their average product rather than their lower marginal product. The competitive wage rate in rural labor markets that is generally reported in the Census would in this case approximate the marginal product of labor and understate the relevant opportunity cost that the individual in family employment considers before migrating. However, if the family invests in the migration of its members, as many studies of migration document (Caldwell, 1968; Hay, 1974) the relevant opportunity cost for family decisionmaking is the shadow price or marginal product of labor. See Sen (1975) for other explanations for dual labor markets and their implications for labor utilization and policy.

TABLE 5

F Tests on Coefficient Restrictions in Migration Model and Relevant Degrees of Freedom

Hypothesis Tested[1] (Regressions Compared in Table 4.	Educational Group			
	No Schooling	Some Primary	Some Secondary	Some Higher
1. With Unrestricted Specification -- Wage Expectation Hypothesis (Regressions 2-1)	.36 (2,364)	.59 (2,365)	5.10** (2,363)	1.44 (2,343)
2. Symmetric Origin and Destination Restriction for U, E and W (Regressions 3-1)	1.17 (3,364)	.56 (3,365)	4.41* (3,363)	2.68 (3,343)
3. With Symmetric Responses -- Wage Expectation Hypothesis (Regressions 4-3)	1.00 (1,367)	1.44 (1,368)	.63 (1,366)	.99 (1,346)
4. Origin Constant Term Effect is negligible (Regression 4)	1.90 (1,368)	.23 (1,369)	1.28 (1,367)	.07 (1,347)

NOTES: Test statistics derived from Table 4.

1 The null hypothesis is that the regression coefficient is zero or that the set of coefficients are equal is rejected at the 5 percent level of confidence (*) or 1 percent level (**), as indicated.

themselves as (openly) unemployed in the urban sector less frequently than would be implied by the HT model.

The symmetry of origin and destination attributes, measured in terms of migration elasticities, is confirmed by the F tests reported in Table 5, row 2 for all but secondary educated, and for both secondary and higher educated the employment rate effects are the only single variable restriction that appears to operate asymmetrically (test not reported). Finally, the statistical significance of the constant term can be interpreted as a test whether nonmigration is a distinct and separable decision, namely, whether $\alpha=\alpha$.* In Table 5, row 5 the comparable F test is reported for whether this constant term effect is different from zero, using regression 4 for comparison purposes. In all education groups the constant is not statistically significantly different from zero. It may still be worthwhile to examine the stepwise migration decisions using a decision tree framework that is consistent with the dichotomous and polytomous logit model applied here.

Predictive Accuracy of Alternative Statistical Models of Migration

The tests of restrictions reported in Table 5 generally support the acceptance of the symmetric uniform logit model as a simple description by one set of parameters of both cross migration rates and the rate of nonmigration. But does the logit formulation represent a notable improvement over the standard "gravity" model as estimated in the unrestricted form of equation (1)? There is no single or simple way to evaluate the goodness of fit across different statistical models based on slightly different bodies of data. I have chosen here to reconstruct the predicted values of all the migration cells based on the logit ordinary least squares estimates of regression 1 in Table 4 and the comparable gravity equation (1) and contrast them on the basis of their coefficient of determination or R^2. First, it should be noted that the gravity and the logit model estimated by OLS (on the odds-ratio) exclude migration cells for which the migration rate is zero; one cell in 380 for men with no schooling, 2 cells for men with secondary schooling, and 22 cells for men with higher education. However, the gravity model also makes no direct use of the 20 nonmigration cells, i.e. m_{ii}'s. This appears to be an advantage to the logit OLS. Comparisons are made, therefore, on both the subsample of observations used for the gravity model (A) and for the sample used by the logit OLS model, plus any remaining zero cells (B). Since my objective is to explain both cross migration and nonmigration in a single model, the sample (B) comparisons are of main interest here.

One may also be concerned that empty migration cells contain information that is disregarded by both of these estimated models, and this information might be of importance, at least for the higher education group where it represents nearly six percent of the sample. To incorporate the zero cells, a nonlinear maximum likelihood (ML) logit estimator is required. Including the mutually exhaustive 20 possible outcomes for each of the 20 regional birth cohorts produced a matrix of outcomes that exceeded the capacity of the available Nerlove-Press logistic program (1973), and the model did not conform to the attribute differenced dimensions required by the Manski-McFadden conditional logit program (nd), even if a problem of this size had been computationally possible. As a second best solution, the 380 observations that excluded the 20 nonmigration cells were used to estimate an unrestricted dichotomous ML logit model at moderate cost. Though the program did not always appear to converge, virtually identical parameter vectors were estimated from various starting values. The omission of the residual m_{ii}'s should not bias the ML logit estimates, but presumably reduces their efficiency.

Table 6 reports the calculated R^2 for the three estimators for both Samples (A) and (B). It is somewhat surprising that the gravity model does quite well in predicting the m_{ii}'s except among the higher educated group; the R^2 actually increases for the gravity model based on the full sample (B) compared with (A), even though it does not explicitly use information from these added sample points. The anticipated deterioration in gravity results for the (B) sample occurs only for the higher educated men, where the R^2 plunges from .54 to .06.

More in accord with expectations, the logit OLS outperforms the gravity model in all education groups for both samples (A) and (B). Clearly these migration data are better fit by the cheaply calculated OLS logit model than by the gravity model, even if only cross migration rates are of interest, i.e. sample (A).

More surprising is the superior performance of the maximum likelihood logit estimators, which outperform the OLS logit in every instance. Even though the ML logit are not based on the nonmigration cells, they predict them admirably, increasing R^2 for those with no schooling from .96 to .99 (the "nonconvergent" case), the secondary educated from .82 to .89, and the higher educated from .70 to .77. In the primary educated group, where there were no zero cells, the LM logit explained .98 of the variance across all of the cells compared with .95 for the OLS logit. In sum, the OLS logit appears to represent a substantial improvement over the gravity model. The moderately more costly nonlinear ML logit estimates, even when nonmigration cells are omitted, provide a still better fit to these Venezuelan lifetime migration data.

Table 6

Comparisons of Coefficients of Determination[a]

Between Actual and Predicted Migration Probabilities

Based on Alternative Models and Estimation Techniques

Education Group		Function Form and Estimation Technique				
Sample Size and Code	Sample Definition	Gravity Model[b]		Logit OLS[c]		Logit ML[d]
		R^2_1	R^2_2	R^2_1	R^2_1	R^2
No Schooling						
379 A	$m_{ij} > 0$ $i \neq j$	-1.23		-.09		.56
400 B	m_{ij}	.92		.96		.99
Some Primary Education						
380 A	$m_{ij} > 0$ $i \neq j$	.49		.56	.57	.71
400 B	m_{ij}	.93		.95	.95	.98
Some Secondary Education						
378 A	$m_{ij} > 0$ $i \neq j$	.60	.62	.71	.72	.87
400 B	m_{ij}	.64	.64	.82	.82	.89
Some Higher Education						
358 A	$m_{ij} > 0$, $i \neq j$	.54	.57	.76	.75	.84
400 B	m_{ij}	.06	.06	.70	.70	.77

Notes: [a]Since the mean of the predicted values need not be the sample mean, given the form and procedure used to estimate the various models, two ways of estimating the R^2 can differ.

$$R_1^2 = 1 - (\Sigma y^2 - 2\Sigma_i y\hat{y} + \Sigma_i \hat{y}^2)/(\Sigma y^2 - (\Sigma y)^2/N)$$

$$R_2^2 = 1 - (\Sigma(y - \bar{y} - (y - \hat{\bar{y}}))^2/(\Sigma(y - \bar{y})^2$$

where y is the actual and $\hat{y}$ the predicted migration probability, $\bar{y}$ is the mean and $\hat{\bar{y}}$ is the predicted mean, and the summations are over the N observations of the sample. When the same value is obtained to 2 digit accuracy, only one result is reported.

Notes to Table 6 continued:

[b]The gravity model is of the form $\log m_{ij} = \alpha + \beta \log X$ and is estimated by ordinary least squares for $m_{ij} > 0$ and $i \neq j$.

[c]The Logit model can be estimated by ordinary least squares in the odds-ratio form, $\log (m_{ij}/m_{ii}) = \alpha + \beta \log X$ for observations where $m_{ij} > 0$, $i \neq j$ and $m_{ii} > 0$. This procedure imposes the adding up restrictions, i.e., $1 = \sum_j m_{ij}$ for all i.

[d]The maximum likelihood logit parameters are the estimates of the logit function $m_{ij} = 1/(1 + \exp -\beta \ln X)$, that maximize the likelihood function given the observations of migration rates for various cross migration cells including zeros, i.e. all m_{ij}, $i \neq j$. This procedure does not explicitly impose the adding up restriction, i.e. $1 = \sum_j m_{ij}$, for all i.

V. Conclusions and Discussion of Findings

Methodological Conclusions

The traditional double logarithmic migration function has proven an adaptable way to explain patterns of interregional migration. But its satisfactory fit to data from several countries should not deter us from replacing it by a more rigorous statistical framework that both uses more information and plausibly restricts the probability of migration to the zero one interval. Recently added to the tools of econometrics, the polytomous logit framework appears suited to studies of interegional migration and socio-economic mobility.

In this investigation of male lifetime internal migration in Venezuela, the linear (ordinary least squares) estimates of the logit model account for more of the variation in migration rates than do the linear estimates of the double logarithmic gravity model. Nonlinear maximum likelihood logit estimates, moreover, prove superior to the cheaper linear logit estimates. Explanatory variables account for more of the variation in migration rates when they are specified in logarithmic form in the logit model, and this also permits direct tests of several useful economic hypotheses regarding migration behavior.

The normalization of migration flows by the origin population size is consistent with these data, except for the higher education group, where the coefficient on origin population, N_i, is statistically significant and negative. As indicated in this paper, the unrestricted inclusion in the migration function of origin and destination population size variables has little basis in theory, and considerable potential for biasing other parameter values by becoming a proxy for omitted persistent factors that account for historic patterns of migration. In this exercise the importance of both origin and destination population size was most noted among the more educated, the class of skilled workers for whom there may indeed be perceived gains from working in a large metropolitan labor market.*

When regional labor market conditions are expressed in logarithmic form, the uniform logistic model implies that origin and destination variables enter in ratio form, an empirical restriction that is found to be consistent with these data. Schooling and population growth variables, however, do not appear to enter in ratio form; the arithmetic form may be more appropriate.

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*For this group alone the Cobb Douglas specification of the migration function may warrent more study, i.e., $\beta_1 \neq 1-\beta_2$. Research is needed to determine under what conditions and for what problems the constant returns to scale restriction on the migration function makes sense. Perhaps the concentration of the less educated groups in certain regions has not reached the same proportions as for the more educated groups, and thus not evoked the nonlinear restraints on migration rates that is implicit in the constant returns to scale Cobb Douglas function.

Empirical Findings

Within the limitations of a single aggregate cross section, the data from the Venezuelan 1961 Census describe plausibly how patterns of internal lifetime migration might respond to economic and demographic forces. Analysis is limited to men, since observed wage and employment rates should be less biased indicators of productive opportunities for men in all regions.* Four education groups are distinguished because it was expected that labor market opportunities and individual migratory response parameters would differ by education. Tabulations for every region documented the tendency for the more educated to migrate more frequently (Table 1) and to be less deterred by distance and more responsive to relative wage and employment differences(Table 4). Within education groups migration is also greater from regions where school enrollment rates are higher. Thus, disaggregation of the population by education appear justified in the study of migration behavior.

Average wage rates at destination are associated with migration within all four education groups; the elasticity of the migration rate with respect to destination wages ranges between 1.4 to 2.9. Measured origin wage rates,

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*Female migration data were also analyzed with results similar to men's. Several differences might be noted, however. First, women tended to respond more strongly (larger elasticities and t statistics) to male wage rates than to female wage rates, particularly among higher educated women. This is consistent with many women migrating with their husband or in search thereof, in which the market earnings of the male dominate the migration decision given their more frequent participation in the market labor force. The expected income hypothesis is accepted only for higher educated women. Based on women's wages and employment rates, the symmetry of origin and destination effects is rejected, but is accepted if based on the respective male variables. The intercepts were virtually zero in all except the higher education group, where the F was 3.17 (1,220) and the intercept was positive. For women the odds ratio in 14 cells for the secondary educated was zero, and for higher educated women 19 infinite (none stayed at birthplace) and 31 zero.

however, are less of a restraining factor on migration, particularly among the unschooled, but operate, nonetheless, in the anticipated direction. Destination wage elasticities, therefore, exceed origin wage elasticities, confirming that the origin wage may proxy the availability of investable funds which facilitate migration. The differential effect by education level of origin and destination economic variables does not confirm the selectivity hypothesis, which implied that origin characteristics should influence predominantly the less educated and destination characteristics should influence the more educated.

The essential feature of the Harris Todaro model of migration and labor factor markets is that inflexibilities in wage rates across labor markets induce compensating variation in employment rates. Holding constant migration rates, the presumed compensating variation between wages and employment levels is not evident in Venezuela among male migrants with less than a secondary education. For these less educated groups in the labor force the traditional wage gap appears to be the predominant determinant of urban labor force growth and interregional migration. This finding can be explained in several ways, but to test these conjectures is beyond the reach of our data. First, the availability of low paid jobs in the traditional urban sector may accomodate less educated migrants upon arrival at destination. The elastic supply of these low paying jobs reduces measured unemployment and lowers the cost of search for modern well-paid jobs. Alternatively, employment levels may have been atypically low during the 1961 urban recession. Particularly for the less educated, employment levels in 1961 might have deviated widely from those prevalent in the prior two decades of urban expansion when much of the observed lifetime migration occurred.*

*The latter explanation was explored with unemployment/employment data derived from the 1950 Census, which also implied little sensitivity of the less educated migrants to the measured level of employment a decade earlier in a more "normal" expansionary phase of the national economy.

But for men with some secondary or higher education the elasticity of migration with respect to employment is greater than that with respect to wage. For these better educated men the Harris-Todaro (1970) framework may be applicable; the potential origin wage (not conditioned on employment rates) is compared with a weighted product of the destination wage rate and the probability of finding an urban job. Holding migration constant for secondary educated, a one percent increase in destination wage rates is offset by a 3.5 percent decrease in employment rates (or a ten times larger relative increase in unemployment rates). For higher educated men a 2.4 percent decrease in employment rates would offset a one percent rise in destination wages. With this steep trade off between employment rates and wage rates, interregional relative variation in wage rates are observed to be smaller among the more educated than among the less educated, or conversely, levels of unemployment observed among various education classes are about the same (Childers, 1974).

Dualism may be an important aspect of the labor market for less educated workers in Venezuela given the large differences across regions in wage levels. But to characterize the trade-off between wage rates and employment at lower education levels will require additional, more refined information on the traditional and modern subsectors, job turnover and the duration of unemployment, hiring practices and evidence of the consequences of wage controls. Published results of the 1961 Census, unfortunately, do not confirm any interplay between employment and wages for the vast majority of less educated Venezuelan workers.

This interpretation of regional integration of labor markets by educational level is consistent with the diminished deterrent effect of distance on the migration of more educated Venezuelans. Also, the size of population at destination attracts notably the more educated, and there is a corresponding tendency for the more educated to be less inclined to move out of the more populous states. These patterns of behavior may be useful for prediction, but should be interpreted with caution as they probably are attributable to the dynamics of the migration process that are not adequately captured in my static framework.

Population growth appears to be a powerful added force affecting the redistribution of the Venezuelan population, but it apparently operates predominantly through its effect on destination choice, and less on the population at origin. The rate of population growth is today beginning to subside in Venezuela (Cerrutti and Kar, 1975), as urban fertility rates decline. But this emerging trend will not immediately slow rural-urban migration. Moreover, the combination of educational, health and family planning services that might evoke a decline in rural fertility that would slow migration may prove costly, and overtax the commitment of many low income countries to rural development.

According to the estimates reported in this paper, a recession such as that which overtook Venezuela in 1961 and increased sharply unemployment would curtail the urban influx of better educated migrants. But increasing measured urban unemployment would do little to dampen the growth in numbers of less educated job seekers so long as the urban-rural wage ratio remains unchanged. Herein is the thorny and perennial problem of wage and incomes policy for a country experiencing rapid but uneven growth. Until means are found to narrow the gap between urban and rural wages for the least educated workers, the reallocation of the population to the cities will continue.

But this conclusion also has another side. The dire prediction that growth of urban employment might become socially counterproductive, as suggested in the context of East Africa by Harris and Todaro (1970), is not confirmed here for Venezuela. With no observed tradeoff on migration between employment and wage rates among the less educated, there is no reason to assume that socially wasteful urban unemployment is due to urban wage rigidities. On the other hand, the potential social costs of unemployment among the better educated requires much more disaggregated analysis to infer the extent to which this unemployment represents idle resources or occurs in conjunction with job search, mostly among young inexperienced workers. Given the above average human, and, I would expect, physical wealth of men in these better educated classes, it is not clear to me that unemployment is a socially inequitable means of clearing inter-regional labor markets when wages are imperfectly flexible. It remains to be seen, however, whether greater flexibility in wage determination might not foster more efficient national utilization of this pool of skilled labor among competing regional employments.

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