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A LIFE CYCLE THEORY OF MIGRATION: WHETHER TO MIGRATE AS A FUNCTION OF CHANGE

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Abstract

The decision to out-migrate is examined in the context of life time utility maximization, where the flow of utility is a function of one's wage rate, probability of being unemployed, and the nonpecuniary characteristics of the location. It is shown that if information about the characteristics of other localities is reasonably good, migration is a response to changes in locational characteristics or in the utility function by which these characteristics are valued. The traditional model of out-migration in which flows are a function of the size of regional wage and unemployment differentials is shown to apply only to new labor force entrants. Because real wage differentials in the United States are no longer so large, the choice of where to live is dominated by idiosyncratic elements--generally nonpecuniary--of an individual's circumstances and preferences. Since these idiosyncratic elements are reasonably stable over time, out-migration of established working men is primarily a function of changes in the characteristics of the locality.

A test of the changes model is conducted in a sample of blue collar and clerical workers who were employed in 1965. Using a logit specification, the decision to move to a noncontiguous state was predicted with personal characteristics and 50 dummies for state of residence in 1965. These dummies were then regressed on the 1964 level and the '64-'69 change in level of unemployment, wage rate, and other economic variables. Our change model of out-migration was supported by the fact that the coefficients on change variables were significantly higher than the coefficients on the corresponding levels variables.

A Life Cycle Theory of Migration: Whether to Migrate as a Function of Change

In most migration research the "whether to migrate" decision is assumed to involve a comparison of the costs of moving to the present value of the benefits and costs of living in alternative areas [Greenwood, 1975b, p. 398]. The measures of benefits and costs used in empirical work are generally wage levels, unemployment rates, and proxies for the attractiveness of the living environment like climate and air pollution. These models are designed to explain why at the end of the period of analysis movers prefer their new location, B, to their previous location, A. They do not, however, satisfactorily explain why B was not preferred at the beginning of the period as well and, therefore, why the decision to migrate was not taken earlier.

Utility theory implies that an individual's choice of where to live (stocks) should depend on the wage and unemployment levels of a locality. Migration theory assumes that changes in the choice of where to live (flows of investment in migration) also depend upon the level of wages and unemployment in the locality. The second proposition is consistent with the first only if individuals not just markets are in disequilibrium. The period of disequilibrium must be quite long for most studies measure migration over a five-year period. The migration literature has given very little systematic attention to how such individual disequilibria may come to exist and persist. In fact, a formal model placing static locational choice considerations into the dynamic framework necessary to predict decisions to change one's location seems to be absent from previously published work.[1]

This paper will develop such a model. We show that realistic characterizations of migration behavior do not require assumptions that individuals are in disequilibrium. Our model is structured by two noncontroversial assumptions:

(1) rationality--a family lives where it prefers to live, (2) utility functions for locations are quite idiosyncratic. The testable restrictions on migration behavior that are unique to this model are produced by two further assumptions: (1) the idiosyncratic elements of people's preferences and circumstances are reasonably stable over time, (2) people are aware of geographic wage and unemployment differentials and the costs of visiting a high wage city to look for a job are small compared to the present value of the cost of living adjusted wage differential [Lansing and Mueller, 1967].

The analysis of the "whether to migrate" decision that emerges from this model requires the division of the population of potential migrants into three life cycle categories. Established working adults are predicted to migrate primarily in response to <u>changes</u> in wage and unemployment levels of communities or <u>changes</u> in the circumstances or preferences of the individual.[2] Retirees should migrate toward locations with low costs of living and an attractive living environment and away from locations with high money wage rates. People looking for their first permanent job after schooling or service in the armed forces will respond to wage and unemployment <u>levels</u> by migrating away from locations with low wages and high unemployment. Thus, the standard levels model of out-migration applies only to decisions made at the beginning of the work cycle. A levels model of out-migration is derivable for prime-age workers only by changing one of the assumptions of the model. One or some combinations of the following assumptions would be needed: (1) there is no stability over time in the idiosyncratic elements of people's utility functions or circumstances that cause people to differ about the relative attractiveness of alternative locations or to face different costs of moving, (2) information about opportunities elsewhere is available to an individual only for a short period of time, and to only a small and shifting proportion of the population at any given time,

(3) when for a large share of the population there are long lags (10 to 20 years) between the creation of a large differential and its perception or the decision to act on it.

In Section I a life cycle model of utility maximizing behavior is used to derive a number of testable restrictions on relations predicting out-migration.[3] In Section II the data used to test the theory--1970 Census 1/100 sample with information about industry, occupation, and state of residence five years ago--is described. Section III presents the results of these tests and our estimates of the out-migration response to changes in a state's relative attractiveness. Section IV discusses the implications of some of the empirical results and makes suggestions for future research.

I. Theory

The Idiosyncratic Nature of Locational Preferences.

This paper explores where the assumption of rational behavior, low information and search costs, and response lags of no more than one year, takes a theory of migration. These assumptions imply that as evaluated by the individual's utility function that prevailed last period, current residence can be presumed to be preferred location. To be specific, we know that the 1965 location of the jth person is preferred in the sense that the utility of any alternative location is smaller than the sum of the present location's utility plus the costs of moving. In other words, the annualized net benefit of moving, D_{ijo}, is negative for all.

$$D_{ijo} = \frac{D^*_{ijo}}{C_j} = \frac{L^*_{ijo}-(L^*_{ojo}+M^*_j)}{C_j} < 0 \text{ for all i-1...n,} \tag{1}$$

where

i = indexes alternative locations to the present one "0",

j = indexes people,

L^*_{ijt} = discounted utility of a life at the "i"th location at time "t" (t = 0 in 1965),

C_j = the ratio of discounted lifetime consumption to annual consumption,

D^*_{ijt} = the utility of the alternative location "i" minus the sum of moving costs and the utility of the present location,

M^* = moving costs.

It will be useful to separate the "j"th person's annualized net benefit of moving to location "i" into its predictable and idiosyncratic elements.

$$D_{ijt} = D^{p}_{ijt} + \varepsilon_{ijt} \quad , \qquad (2)$$

where

D^{p}_{ijt} is the predicted net benefit of moving to i. The model making this prediction is limited to a small set of measurable locational and personal characteristics as right hand side variables. Except for certain specified interactions between stage in the life cycle and locational characteristics, the parameters of this model are assumed identical across individuals.

ε_{ijt} is the error in this model's ability to predict the net benefit of moving, D_{ijt}. ε_{ijt} captures the idiosyncratic elements of utility functions or circumstances that cause people to differ about the relative attractiveness of alternative locations or to face different costs of moving. ε_{ijt} is assumed to be generated by an autoregressive process $\varepsilon_{ijt} = \rho\varepsilon_{ijt-1} + v_{ijt}$ where ρ measures the stability of the idiosyncratic elements in locational preferences.

Money wage differentials across states are no longer very large.[4] Real wage differentials are even smaller. Consequently, idiosyncratic elements--often nonpecuniary--of an individual's circumstances and preferences often dominate the choice of where to live. A variety of powerful factors produce this diversity:

a. Desire to live near or far from relatives.

b. Unique characteristics of the geographic environment of one's home.

c. Valued friendships that cannot be replaced in the new location.

d. Desire to live within a particular ethnic community.

e. An occupation that involves developing a clientele over time.

f. A good job with a local firm.

Most of these causes of diversity are quite stable over time. Variations across time in the value of ε_{ijt} seem to reflect events that are both uncommon and unpredictable.

a. Divorce, separation, death in the family.

b. Flood, tornado, or fire.

c. Friends dying or moving away.

d. Urban renewal.

e. Bankruptcy, disbarment.

f. Loss of job, plant closing.

Consequently, the autocorrelation parameter ρ is expected to be significantly greater than zero. While not equal to one, it is expected to be closer to one than zero. If the assumption of low costs of information were to be dropped, the ε_{ijt} would reflect the variance in information as well as the variance of idiosyncratic preferences and ρ would capture the stability of this information as it currently captures the relative stability of idiosyncratic preferences.

The Decision to Migrate

Migration is a decision to change one's location. If at the beginning of the period one's current location is the preferred location, out-migration will occur only if changes occur in either one's circumstances, the environment, or in the valuations placed upon environmental characteristics. The individuals will decide to leave his 1965 location if and only if D_{ij1} is positive for some "i".

$$P(OM) = P(D_{ij1} > 0, \text{ for some } i; D_{ijo} < 0, \text{ for all } i). \tag{3}$$

Let us rewrite D_{ij1} as a function of the sum of the beginning of the period level of D_{ijo} and its change over the period in which migration is measured.

$$D_{ij1} = (D^P_{ij1} - D^P_{ijo}) + \varepsilon_{ij1} - \varepsilon_{ijo} + D_{ijo}$$

$$= \frac{dD^P_{ij}}{dt} + (\rho-1)\varepsilon_{ijo} + v_{ij1} + D_{ijo}$$

$$= \frac{dD^P_{ij}}{dt} + (1-\rho)(D^P_{ijo} - D_{ijo}) + v_{ij1} + D_{ijo}$$

$$D_{ij1} = \frac{dD^P_{ij}}{dt} + (1-\rho)D^P_{ijo} + v_{ij1} + \rho D_{ijo} \quad . \qquad (4)$$

Since our data will be persons who live in location "0" in 1965, our sample is selected on the value of D_{ijo} ($D_{ijo} < 0$ for all i). This conditional distribution of D_{ijo} is assumed to be independent of the measurable characteristics of the location and is, therefore, a part of the model's disturbance.

The degree to which idiosyncratic elements of D_{ijo} are stable over time is a central element of this migration model. As measured by ρ it determines the extent to which beginning of period levels of a location's characteristics (D^P_{ijo}) enter the model.

The first two terms of (4) represent the predictable element of the "whether to migrate" calculation. Referring back to equation (1), we see that D_{ijo} and consequently D^P_{ijo} have two parts: a comparison of the desirability of living in the alternative location $(L^*_{ijo} - L^*_{ojo})/C_j$ and an annualized moving cost figure (M^*_j/C_j). The costs of moving $\left(M^*_j\right)$ are both monetary and psychological: transportation for family and furniture, search costs for housing and a job, costs of selling an owner-occupied home, unhappiness in the lonely period before new friends are found, and the disruption of children's schooling. They are "once and for all" costs. Unlike the other elements in a family's calculation (differences in wage levels and climate, etc.), total moving costs do not depend on how long the family expects to stay in the new location. As a consequence, when annualized (divided by C_j), they rise with age or a shortening of the planning horizon. Differentiating annualized moving costs, M_j, with respect to time:

$$\frac{d\,M_j}{dt} = \frac{d}{dt}\left(\frac{M_j^*}{C_j}\right) = \frac{\frac{dM_j^*}{dt}}{C_j} - \frac{M_j^*}{C_j^2}\,\frac{dC_j}{dt} \quad . \tag{5}$$

Annualized moving costs (M_j) rise with age both because their total increases with age or tenure in one place [the first term of (5)] and because the period over which these investment costs must be recouped shortens (the second term). People with initially high moving costs (M_j^*) have faster rising annualized moving costs as well. This is part of the reason why people with high moving costs are less likely to migrate.

The Utility Function for Locations

The theory developed so far suggests that most decisions to migrate are a response to <u>changes</u> in the perceived relative attractiveness of one's location. Most of the time the change is in what is viewed as important, not in the actual character of the metropolitan area. Often the changes of preferences occur in an essentially random pattern. Even when the change of preferences is caused by a change of circumstances, it is often unpredictable. The predictable shifts of preferences for locations are associated with watersheds of the life cycle: entering the civilian labor force for the first time upon leaving school or the armed forces and retiring. At the time he decides to look for a civilian job, the location of a school leaver or army dischargee is generally a consequence of birth or army assignment or a choice based primarily on the availability of advanced schooling. His current location carries very little information about the type of place that individual considers desirable as a place to work and live in for the rest of his life. For new labor market entrants, we would expect that migration would be away from locations with low wages, high prices, and high unemployment. Their out-migration decisions would, therefore, be consistent with the traditional economic model of migration in which the decision to leave is a function of relative real wages and relative unemployment rates.

Most migrants, however, are not new labor force entrants. Of the 4.6 million employed males 25 years of age or over who had moved across state lines between 1965 and 1970 only 680,000 were leaving the armed forces and only 1.8 million did not work in 1965 or did not report what their work was. People who are employed at the beginning of the period over which migration is measured must be presumed to have already made a choice of where to live, taking into account relative levels of unemployment and wages. For people who retire during the period, the importance of the employment characteristics of location will decline precipitously and the relative importance of quality of life characteristics (weather, air pollution, and cost of living) will rise. Retirees will, therefore, tend to migrate away from places with high wages and low unemployment and toward places with low costs of living. For the great bulk of workers whose labor force commitment remains essentially constant, the fact that the individual chose to move into or remain in a given state is a good indication of his preferences.

We will now derive a more specific characterization of an out-migration function from a Taylor approximation of an arbitrary utility function. The utility function has the following characteristics.

(a) Because moving costs rise over time, the individual or family evaluates alternative locations on the assumption that they will remain there the rest of their lives. Consequently, the model does not apply to young people moving to another state to attend school.

(b) Future consumption benefits are discounted over one's entire lifetime, while future employment benefits are discounted over one's working lifetime.

(c) Rates of change in the relative attractiveness of a particular location are not extrapolated into the future, and to the extent that existing differentials are expected to diminish in the future we may capture this by placing a risk premium on the discount rate.

(d) Except for the changing relative importance of work and of consumption-related characteristics of a location, the individual expects his utility function, including the effects of unmeasured qualities of the location, to be stable over time.

We work with a utility function defined over annualized present values of locational characteristics. This together with the choice of scale for the utility index is designed to maintain approximate stationarity of the disturbance.

$$L^p_{ijt} = g(r_{jt}W_{it}, A_{it}, Q_{it}) , \tag{6}$$

where

t indexes the date (units of time are five years and 1965 is year "0"),

L^p_{ijt} = the predictable element in the annualized utility of living in location i for the jth person,

r_{jt} = the ratio of the present value of a dollar discounted over the working lifetime to the present value of a dollar discounted over the entire lifetime,

W_{it} = a vector of employment related characteristics of the "i"th location in the "t" time period,

Q_i = a vector of stable quality of life characteristics, and

A_{it} = quality of life characteristics that can change over time.

Let us take a first order Taylor expansion plus interaction terms of this utility function around its value for location zero in 1965 (L^P_{oo}). By treating W, A, and Q as variables, we may use this function to evaluate the disadvantages of location i.

$$L^P_{ijo} - L^P_{ojo} = \beta r_{jo}(W_{io}-W_{oo}) + \alpha(A_{io}-A_{oo}) + \gamma(Q_i-Q_o) \quad , \qquad (7)$$

where

β, α, and γ are vectors of first derivatives of the utility function with respect to locational characteristics.

Multiplied by $(1-\rho)$ expressions like (7) are a component of the second term of migration function (4).

The first term of (4) measures the change in attitude toward locations over time. By treating W, A, r_j, and α as variable, we may evaluate the impact of these changes in preferences or community environment on the utility derived from a location. For one's current location:

$$L^P_{oj1} - L^P_{ojo} = \frac{dL^P_{oj}}{dt} = \beta r_{j1}\frac{dW_{o1}}{dt} + \beta W_{oo}\frac{dr_j}{dt} + \alpha\frac{dA_{o1}}{dt} + A_{o1}\frac{d\alpha}{dt} \; . \qquad (8)$$

An economic model of migration must condition its structure on the stage of the individual in the life cycle. The effect of the life cycle on the evaluation of the relative attractiveness of alternative locations is captured by r_j, the ratio of the present value of a dollar received over a working lifetime to a dollar received over one's entire lifetime. Two terms involving r_j enter (8). In the first term the value of r_j at the end of the period is a multiplier of the change in employment-related characteristics over the period. In the second term the change in r_j over the period is a multiplier of wage and unemployment levels at the beginning of the period. Individuals, whose planning horizon (period over which he plans to remain in one location) does not include civilian employment, have an r_j of zero. Thus, a full-time student with no part-time job, whose planning period ends with graduation, will have $r_{jo} = 0$. When he graduates and enters the labor market, his r_{j1} will rise to .96. Since $r_{j1} = \frac{dr_j}{dt} = .96$, the first two terms of (8) simplify to $.96\beta W_{o1}$. Thus, for new labor force entrants migration is a function of end of period levels of location characteristics.

For people who are working at the beginning of the period, r_j and the pattern of changes in r_j is very different. Aging results in actual retirement or disability for some and the closer approach of retirement for others. Consequently, work-related characteristics of a location decline in importance relative to quality of life characteristics. The decline of r_j from .95 at age 25, to .84 at age 45, and .54 at age 60 reflects this life cycle shift.[5] For prime-age workers the rate of decline is not substantial, however, (over the previous five-year period dr_j/dt is $-.01$ at age 25 and $-.04$ at age 45) so the second term of equation (8) is of minor practical importance. For this group it is changes in the employment characteristics of the location [the first term of (8)] that are most important.

Only when retirement nears or occurs does the second term of (8) regain its importance. A recently retired 65-year-old has a dr_j/dt of $-.54$ and a r_{j1} of zero. A working 65-year-old has a dr_j/dt of $-.105$ and a r_{j1} of .44. The substantial negative values for dr_j/dt meant that age brackets in which large proportions of the population are retiring should tend to migrate away from cities and states with high wage levels and toward localities with low costs of living and an attractive quality of life.

The last two terms of equation (8) capture the effect of shifts in the quality of life (e.g., the introduction of air conditioning, reductions in crime, improvements in local public services) and shifts in views about the relative importance of these characteristics.

Empirical Specification of the "Whether to Migrate" Decision

All the ingredients for analysis are now available. By substituting (5), (7), and (8) into (4), we derive a testable parameterization of the net benefit calculation (D_{ij1}) in terms of observable characteristics of the family and locations. The empirical testing of this model will focus on a dichotomous choice, the "whether to migrate" decision, so further simplification is possible. The individual will move from his current state of residence if any of the D_{ij1} are positive, or stated another way, if, D_{mj1}, the maximum D_{ij1} is positive. When (4) is rewritten for an unknown D_{mj1}, a number of things change: the comparison location has national mean characteristics and the disturbance ($v_{mj1} + \rho D_{mjo}$) now captures systematic changes over time in the measurable characteristics of the best alternative location ($L_{mj1} - L_{mjo}$) and the residual effect of beginning of period levels of this location [$(1-\rho)L_{mjo}$]. The logit distribution is chosen as the specification for our disturbance ($v_{mji} + \rho D_{mjo}$). For the analysis of dichotomous choice, a logit specification has a number of advantages: it has a computable likelihood function and interpretable coefficients. The most common alternative to Logit analysis is OLS on a linear probability function. Linear probability models have the disadvantage of not constraining a probability to the zero-one interval. When estimated on individual data, they suffer from heteroskedascity. After all the substitutions are made the predictable element of the net benefit of moving may be written:

$$D^p_{mj1} - \rho D^p_{mjo} = -\beta_1 r_j \frac{dW_{o1}}{dt} - \alpha \frac{dA_{o1}}{dt} \quad \text{(impact of changes in environment)} \tag{9}$$

$$-\beta W_{o1} \frac{dr_j}{dt} \quad \text{(decline in relative importance of employment characteristics due to aging)}$$

$$-A_{o1} \frac{d\alpha}{dt} \quad \text{(impact of changes in tastes)}$$

$$-(1-\rho)(\beta W_{o1} + \alpha A_{o1} + \gamma Q_{o1} + M_j) \quad \text{(continuing importance of levels due to transitory disturbances)}$$

$$-\frac{1}{C_j}\frac{dM^*_j}{dt} + \frac{M^*_j}{C_j^2}\frac{dC_j}{dt} \quad \text{(rising annualized moving costs).}$$

Given the sample is limited to those with $D_{mjo} < 0$, the logistic error specification implies:

$$\text{Prob}\ (D^p_{mji} \geq v_{mji} + \rho D_{mjo}) = \frac{1}{1 + e^{-b(D^p_{mj1} - \rho D^p_{mjo})}} \quad \text{or alternatively} \tag{10}$$

$$\text{Log}\ \frac{\text{Prob(OM)}}{1 - \text{Prob(OM)}} = b(D^p_{mj1} - \rho D^p_{mjo}) \tag{11}$$

II. Data and Methodology

Data on state of residence changes between 1965 and 1970 by people who were employed in 1965 were used to test the theory outlined above. The first step was to use individual data from the 1/100 Public Use sample of the 1970 Census to estimate a logistic model of out-migration for separate broad occupational categories. Individual rather than aggregate data was chosen because the logistic model of individual behavior that can be derived from the theoretical model does not aggregate to a simple logistic function when the probabilities of individuals migrating are added together. Furthermore, migration models tested in aggregate data often suffer from a multicolinearity problem. When states are the observations, only 51 observations are available. Controls for individual level determinants of migration--birth in another state, education, marital status, age, race, and industry or employment--must often enter the regression equation as state or SMSA averages. These variables are often colinear with the environmental variables--wage level, climate, price level, the unemployment rate--that are necessarily defined at the aggregate level. Models estimated on individual data are necessarily more efficient estimators of the individual determinants of migration and, therefore, partition the variance between individual and environmental factors more effectively.

The primary disadvantage of estimating migration models on individual data is that if the correct functional form is used (a logistic or probit specification), an iterative estimation procedure that must evaluate the ability of the model to predict each observation's behavior at each step must be used. Each model specification requires a separate run. When some right hand side variables are continuous and the number of observations is large (as they must be to obtain stable estimates of behavior when the event being predicted is very infrequent), this research strategy can become very expensive. If all right hand side variables can be made categorical, however, much more efficient techniques for fitting logistic models become available. The environment was made categorical by treating each state as a separate category. The following characteristics of the individual were controlled: age (in seven categories), having been born in one's 1965 state of residence, marital status in 1965, race and whether ones industry of employment in 1965 had more than 35 percent of its workers in establishments employing more than 250 people. Models were estimated for three separate broad occupations: male craftsmen, male operatives and laborers, and clerks and retail sales workers of both sexes. Service workers, managers, professional and technical and nonretail sales workers were excluded because data on wage rates for these occupations were not available by state for noncensus years. College students, people employed in agriculture or by the federal government in 1965, and people over the age of 59 were excluded from the sample.

Not all moves across state boundaries involve a change of jobs or a change of labor market. Movements to an adjacent state were considered to be moves within a given labor market when the ratio of the number of people crossing the state border in either direction to work to the number of migrants between the two states is greater than .5 (note that the number of household units migrating will typically be between one-half and one-quarte of the number of migrants).

Table 1 presents measures of the impacts of personal characteristics on out-migration. The logit coefficients are the change in the log odds of out-migration due to the characteristic. The probabilities are very small, so that these coefficients closely approximate those that would have been obtained if ln P had been the dependent variable. The anti logs of the tabulated coefficients give the multiple by which the odds of out-migration change due to having the characteristic specified. For instance, the odds a laborer or operative will migrate are 3.42 times greater if the individual was born in another state. Since we are predicting a very unlikely event, the impact of a variable on the odds closely approximates its effect on probability. If an operative born in the state were to have a 4 percent probability of migrating, another otherwise identical individual born outside the state would have a 13 percent probability of migrating. The multiple for probability is only slightly smaller than the multiple for odds.

A χ^2 test of the significance of adding the environmental dimension (the 50 categories for state) to the model was performed. The reduction in χ^2 obtained when main effects for each of the 50 states was added to the model ranged between 221 and 254. Under the null hypothesis that the restricted model is the true model, this statistic is distributed as χ^2 with 50 degrees of freedom. The critical χ^2 for a .005 significance level is 79.5, so we reject the null hypothesis of no environmental effects.

Table 1. Logit Model of Out-Migration of Workers between 1965 and 1970

	Laborers and Operatives (N=26579)		Craftsmen (N=21,474)		Clerks inc. Retail Sales (N=26,608)	
	logit coef.	anti log of coef.	logit coef.	anti log of coef.	logit coef.	anti log of coef.
Born different state	1.23	(3.42)	1.30	(3.68)	1.21	(3.35)
Not married in 1965	.21	(1.23)	.26	(1.30)	.01	(1.01)
Industry plants small	.25	(1.28)	.24	(1.27)	.10	(1.105)
Black	-.40	(.67)	-.21	(.81)	-.43	(.65)
Age in 1965						
14-17	.85	(2.34)	.59	(1.81)	.88	(2.42)
18-22	.66	(1.94)	.52	(1.69)	.70	(2.02)
23-27	.30	(1.36)	.30	(1.35)	.11	(1.11)
28-32	.03	(1.03)	-.06	(.94)	-.19	(.82)
33-39	-.68	(.51)	-.42	(.66)	-.55	(.58)
40-49	-.87	(.42)	-.59	(.55)	-.74	(.48)
50-59	-.30	(.74)	-.34	(.71)	-.26	(.77)
Chi-square test of the addition of state environments						
χ^2 of Model without States	2390		2062		2288	
χ^2 of Model with States	2146		1841		2034	
Reduction of χ^2	244		221		254	

III. Empirical Tests of the Change Model

The second step of the analysis was to regress the coefficients of the state dummies on variables parameterizing the state's economic environment. The accuracy with which it is possible to estimate the effect parameter for each state varies, however, so a weighted least squares technique is required. The variance of our measurement error is inversely proportional to the number of workers in the occupation in that state. The range is substantial--from 37 to 2996. Since the logit becomes increasingly sensitive to changes in $\hat{P}$ as P approaches 0 or 1, measurement error also depends on P. When the data is based on random sampling from a binomial population, the asymptotic variance of the estimate of a logit is $V^m_{(iL_i)} = 1 / N P_i(1-P_i)$ [Theil, 1971: 636].

The second source of error in our model predicting the coefficients on the state dummies is equation error. If the true model has a logit functional form, this error is constant. We chose to assume, a priori, that when a state is very large (N = 2000) and has the national average out-migration rate (P = .05), the equation error will have a variance four times the measurement error. This implies that in smaller states of 500 or so (eg., operatives in Alabama) equation and measurement error are approximately equal and that in the smallest states (Nevada, the Dakotas, Wyoming, Vermont, etc.) measurement error is ten times equation error. It is possible to check the validity of our weighting assumption by comparing the residual variance implied by it to the residual variances corrected for degrees of freedom of the estimated regressions.[6] Our weighting scheme does quite well for the implied residual variances--.114 for operatives, .116 for craftsmen and .099 for clerks--and are quite close to the residual variances of our models (last column of Table 2).

Traditional migration models specify out-migration as a response to wage and unemployment differentials across geographic areas. The typical empirical specification involves making out-migration a function of an area's average unemployment rate and wage rate during the period of analysis. This specification, however, confounds the effects of the level of wages or unemployment with the effect of change in a community's wage or unemployment rate relative to other communities. The separate effect of levels and changes may be distinguished by defining two variables: the level of the characteristic at the beginning of the period and its change over the period. Wage and unemployment rate differentials across communities are remarkably stable over time. In state data two-year averages of manufacturing wage rates four years apart have a .98 correlation. Two-year averages of state unemployment rates five years apart (1963-1964 vs. 1968-1969) have a .80 correlation. Consequently, a good test of the pure level's hypothesis is a regression of out-migration state dummies on 1964-1965 wages and unemployment rates.[7]

The pure levels hypothesis fails this test miserably. Competing with two variables that control for the size and urbanization of the state,[8] the two measures of the local economic environment in 1964--the wage and unemployment rate--are insignificant in every case (see Table 2). For clerks they have the wrong sign.

The pure changes hypothesis (line 2 of each panel) is more successful. The wage change variable is significant in all three regressions and the $\bar{R}^2$ rises substantially.

Entering both changes and levels of the economic characteristics does best of all. All four variables become significant with the correct sign in the operatives equation. In the equation for craftsmen both change variables are significant. The coefficients of the levels variables are almost identical

Table 2. The Effect of the Economic Environment on Log Odds of Out-migration

		Manufacturing Wage Level	Manufacturing Wage Change	Unemployment Level	Unemployment Change	High Wage Percent 1965	Per Pupil School Expend.	Heating Degree Days	Log Employ.	Log SMSA Pop.	$\bar{R}^2$	Resid. Variance
Male Laborers and Operatives	Pure Levels	-.09		2.5					-.247	-.102	.470	.159
		(.21)		(1.02)					(2.34)	(1.36)		
	Pure Changes		-4.97		8.5				-.124	-.184	.529	.140
			(2.54)		(.93)				(1.16)	(2.77)		
	Levels and Changes	-1.23	-7.65	12.9	30.5				-.094	-.126	.598	.120
		(2.53)	(3.37)	(2.07)	(2.49)				(.94)	(1.91)		
	Full	-.60	-7.05	8.6	21.2	-2.59	-1.52	.046	-.061	.001	.705	.089
		(1.14)	(3.48)	(1.51)	(1.89)	(2.64)	(3.52)	(1.31)	(.62)	(.02)		
Male Craftsmen	Pure Levels	.32		6.0					-.244	-.11	.412	.193
		(.66)		(1.09)					(2.13)	(1.35)		
	Pure Changes		-6.34		8.5				-.117	.176	.492	.167
			(3.00)		(.82)				(1.01)	(2.46)		
	Levels and Changes	-.775	-7.67	13.6	30.0				-.091	-.141	.522	.157
		(1.40)	(3.04)	(1.86)	(2.06)				(.80)	(1.87)		
	Full	-.35	-6.81	12.6	22.2	-3.92	-1.19	.100	-.044	-.042	.626	.123
		(.57)	(2.91)	(1.85)	(1.62)	(3.45)	(2.40)	(2.48)	(.39)	(.50)		
Clerks	Pure Levels	.39		-2.2					.264	-.102	.562	.112
		(1.04)		(.51)					(3.44)	(1.87)		
	Pure Changes		-3.47		9.6				-.223	-.106	.608	.100
			(2.30)		(1.24)				2.94	2.22		
	Levels and Changes	-.23	-4.08	-.7	9.7				-.227	-.094	.593	.104
		(.52)	(2.14)	(.12)	(.87)				(2.89)	(1.79)		
	Fall	-.06	-3.35	-2.1	3.5	-1.96	-.46	-.044	-.155	-.064	.605	.100
		(.11)	(1.73)	(.35)	(.30)	(1.93)	(1.07)	(1.24)	(1.70)	(.90)		

to the corresponding coefficients in the operatives. Despite this similarity of results, however, these levels coefficient equations are not statistically significant.

The size and significance of the change variables in the blue collar regressions is doubly impressive when it is realized that it occurs in the face of a potential simultaneity problem that will, if it exists, bias the coefficient toward zero. The model specifies that out-migration responds positively to a rise in the unemployment rate and negatively to a rise in the state's manufacturing wage rate. Causation could also go the other way, however. By tightening the local labor market, an exogenous increase in out-migration might cause wage rates to rise and unemployment rates to fall. Instead of the negative association between out-migration and wage rises of our migration model, this phenomenon would tend to produce a positive association. A negative feedback simultaneity of this type biases coefficients on the endogenous variables toward zero. If the simultaneity is really powerful, the sign of the estimated relationship may be reversed.

The models presented in this paper are not likely to suffer from this problem, however. As the residual variance of the state level out-migration relationship approaches zero, the potential for simultaneous equations bias approaches zero [Kadane, 1971]. The three most important determinants of local out-migration rates, age structure, the proportion of population born outside the state, and employment status at the beginning of the period were effectively controlled for by either the selection of the sample or in the individual level model of the first stage of analysis.

While the two-stage approach to estimating an out-migration relationship tends to increase measurement error problems, it minimizes equation error.

Only equation error can produce a simultaneity bias. Relative to the overall across state variance of employment growth rates, the equation error of the state level out-migration relationships is quite small.[9]

In our models for clerks, the levels variables and the change in the unemployment rate are consistently insignificant. Only the growth of the manufacturing wage rate is significant. The small and insignificant coefficients could reflect a much lower responsiveness of this occupation's migration decisions to economic incentives or a number of specification problems. Women are over 70 percent of the clerical sample. Their migration behavior may be responding to their husband's job opportunities and since many of these men will have white collar occupations, the manufacturing wage may be a poor proxy for their opportunities. If a larger sample could be obtained it would be desirable to model the behavior of men, married women, and single women separately and to use levels and changes of clerical wages as measures of the economic environment. The contrast between our blue collar and white collar results suggests that it is important for the wage rate measure used to apply to the population studied. Support for this observation comes from the fact that when per capita income growth is substituted in the blue collar regressions for the growth of manufacturing wage rates, the explanatory power of the model deteriorates markedly.

Changes in manufacturing wage rates and unemployment rates (relative to other states) have substantial impacts on the propensity of male blue collar workers to leave a community. The coefficient on wage change in the third equation of the first panel of Table 2 implies that a two standard deviation (.07) fall in a state's wage rate relative to other states will have a once and for all effect of increasing the log odds of out-migration over the five-year period by .54 (i.e. 7.65 x .070). This translates into an increase in probability of an individual migrating during the period of the fall in local wage rates of approximately 60 percent. If the state's wages do not continue to fall relative to the rest of the nation, propensities to migrate will in later time periods almost completely return to their former level. The levels coefficient implies that a permanent fall in wage rates of 7 percent will, after the initial impact is over, have a continuing tendency to raise the log odds of out-migration by .086 percent.

While impacts on log odds of out-migration are large, the implied reduction in population is small. A permanent fall in a state's relative wage of 7 percent will induce a rise in the state's blue collar out-migration rate from .0426 to .0706 in the first five-year period and a rise to .046 in later periods.

A permanent rise in a state's unemployment rate relative to other states of 2 percent will increase the log odds of out-migration by .61 during the period of the rise and will raise the log odds by .258 in later periods. The typical individual's probability of out-migration will be 77 percent higher in the initial period and remain 28 percent higher in later periods as long as the new unemployment differential is maintained.

Here again the implied population reduction is small. Out-migration is predicted to rise from 4.26 percent to 7.56 percent in the first five-year period and to 5.47 percent in later periods. Only when a real disaster strikes and a state experiences a simultaneous 7 percent decline in its relative wage and a 2 point rise in unemployment does the resulting population reduction become substantial. Under these circumstances out-migration rises to 12.08 percent during the first five-year period and to 5.9 percent in later periods.

A number of tests of the sensitivity of our results to specification were made. Adding a third measure of the pressure of labor demand--employment growth rates--did not appreciably change the coefficients on the wage and unemployment variables.[10]

The only specification change that had substantial effects on our estimates of wage and unemployment effects was adding another measure of the attractiveness of local job opportunities (see line 4 of each panel). When the ratio of employment in high wage manufacturing industries (all manufacturing minus lumber, textiles, apparel and footwear) to total state employment is added, our estimate of the wage level's impact on out-migration falls. This is as expected. Less expected is the one-third decline in the coefficients on the unemployment rate and its change. The reason for this is that 1965 to 1970 was a period of rapid expansion by the high wage manufacturing sector. In states with a large high wage sector, blue collar workers who wanted to switch into the higher wage sector would have had the opportunity to do so without migrating. In states with small high wage sectors, migration was necessary to obtain a job in the high wage sector. States with large high wage manufacturing sectors also happened to have larger than average drops in unemployment ($r = -.38$), so adding this variable lowers the coefficients on the change in unemployment. Note that while the proportion in the high wage sector is technically a levels variable, it is here interpreted as a proxy for change in job opportunities. If our interpretation is correct, the effect of this variable should disappear during periods of declining labor demand by this sector.

For the models that include comparably defined change and level variables, it is possible to derive estimates of ρ, the relative importance of stable idiosyncratic elements of the individuals utility function for location (see Table 3). The estimates of ρ are generally in the neighborhood of .8 when derived from the wage coefficients and generally about .5 when derived from the unemployment coefficients.

Table 3. Estimates of ρ: the Autocoorelation over a five-year period of Idiosyncratic Elements in the Utility Function for Locations

(In parenthesis is the t statistic of the hypothesis that $\hat{\rho} > 0$.)

	Manufacturing Wage Rate			Unemployment Rate		
	Me3	M4[a]	M4 w/o High Wage %	M3	M4	M4 w/o High Wage
Operators	.79 (3.14)	.86 (3.63)	.83 (3.74)	.53 (1.89)	.54 (1.51)	.61 (2.09)
Craftsmen	.85 (3.04)	.90 (3.16)	.83 (3.23)	.50 (1.47)	.38 (.90)	.51 (1.65)
Clerks	.895 (2.26)	.17 (2.01)	.88 (2.24)	1.02 (1.26)	.55 (.66)	1.03 (1.26)

*These estimates of ρ assume that the positive skewness of the D_{10} distribution is independent of the beginning of period levels of economic characteristics of the location and that $dr_j/r_j = -.05$. The ratio of $\beta_1/\beta_\Delta = (1-\rho) -.05$ so $\hat{\rho} = .95 - (\beta_1/\beta_\Delta)$. If the positive skewness of D_{mjo} is positively related to high wages and low unemployment rates, our estimates of $\hat{\rho}$ are too low. The test for $\hat{\rho} \not> 0$ was β_Δ of hypothesized sign and $|\ .95\ \hat{\beta}_\Delta - \hat{\beta}_1\ | \not> 0$. Since it is a one tail test the critical t for a .05 significance level was 1.68.

[a]These estimates are biased up by the tendency of the proportion high wage variable to pick up part of the effect of wage levels on out-migration.

A test of whether ρ is greater than zero may be equivalently framed as a test of whether the change and level coefficients for wages and unemployment are equal and of the correct sign. The $\hat{\rho}$'s derived from the wage variables are always significantly greater than zero. The $\hat{\rho}$ derived from the unemployment variables have the correct sign, but only a third of them are significantly different from zero.

The reason for the discrepancy between the two estimates of $\hat{\rho}$ is that the statewide unemployment rate does not directly enter the utility function. What enters the utility function is one's expected experience with unemployment. When a worker becomes unemployed, it is natural for him to revise his expectations about future unemployment and to revise his assessment of the costs of migrating. Consequently, a higher unemployment rate in a particular state has both "change" and "level" effects. The levels effect is that the employed generally expect to experience more unemployment over their life time. The change effect is that more people are entering spells of unemployment for the first time. Consequently, the statewide unemployment rate and its change do not really fit the neat level versus change typology developed in the theoretical section.

IV. Conclusions and Possible Extensions

It has been shown that when costs of information about alternative locations are low, lifetime utility maximization yields very different specifications of the migration relation at each stage of the life cycle. People looking for their first permanent job should base their migration decisions on the size of wage and unemployment differentials as is conventionally hypothesized. People who are retiring, however, should be migrating away from high wage areas toward places with low costs of living and attractive quality of life characteristics. Prime-age working people should migrate primarily in response to changes in the environment, their circumstances or their preferences. This implication of the life cycle theory was tested against the conventional levels specification in a large sample of clerical and blue collar workers in 1965. An efficient technique for fitting logit models to large bodies of micro-data was developed for this purpose. The life cycle model of migration was supported by the discovery that changes in wage and unemployment rates had much stronger impacts on out-migration than the levels of these variables.

Our results also have important implications for other areas of economics. A frequent assumption in open models of subnational economies is that migration is costless and quickly responsive to real wage differentials and that all individuals have the same locational tastes. [Evans, 1972, Yinger and Danziger, 1976]. The last assumption can be interpreted in the context of our model as an assumption that the variance of $(v_{mpl} + \rho D_{mjo})$ is very small and that consequently coefficients on changes in wage and unemployment rates should be quite large.

Despite the fact that this study finds out-migration much more responsive to changes in wage levels than other studies, the typical state's elasticity of labor supply due to out-migration is only .4. Even if we add an in-migration response and adjust for the likely higher responsiveness of new labor force entrants the implied aggregate five-year labor supply elasticity for declines in relative wages cannot be much above one. Out-migration responses of this magnitude suggest that labor supply responses alone are not sufficient to maintain the real (adjusted for housing and living cost differentials) wage level of a city that loses one of its major employers.

A further implication of our rejection of the identical utility function assumption is that hedonic cross-section wage regressions do not identify the structure of demand for the characteristics of localities [Rosen, 1974].

Hopefully, the success of this preliminary attempt to build a formal dynamic theory of migration will encourage others to develop even more realistic theories. It would be interesting to see what happens when one drops the assumption that locations are evaluated as if the family expects to remain there. The probabilistic nature of unemployment and of job search could also be explicitly modeled. We attempted to calculate the distribution of the error term ($v_{mji} + \rho D_{mjo}$) from assumptions about the distribution of its components but obtained a computationally intractable result. Possibly others will succeed where we have failed. The explicit dynamic modeling of the "where to migrate" decision should be attempted. Movements within an SMSA may also be modeled in a similar manner.

This paper has tested only one of the implications of the life cycle theory. Many others remain to be tested. A comparison of empirical out-migration models for retirees, new labor force entrants, and prime-age workers will provide a powerful test. Another test is possible by comparing migration models of families at different points in the child-rearing cycle. Families with young children should be migrating toward cities that are good places to rear children and that have quality schools. Once their children have left the home, migration should tend to be away from locations that are particularly attractive for bringing up children.

Implied in the empirical methodology we have chosen is the view that only in micro data will it be possible to satisfactorily model out-migration behavior. The success of the two stage procedure for separating the effects of location from the effects of individual characteristics will hopefully recommend its use to others. The contrast between our blue and white collar results should remind us that high wages and shortages in one occupation's labor market may be balanced by a surplus for another occupation. Occupational disaggregation and measures of the local economic environment specific to that occupation are, therefore, highly desirable.

NOTES

[1]Lowell Galloway [1967], Interindustry Labor Mobility in the United States 1957 to 1960, for instance, does use a utility function to motivate his theoretical model but before the choice of industry decision is taken up all locational characteristics have been aggregated into a shadow wage. Once shadow wages are derived migration, is assumed to be a simple function of the shadow wage differential. See also Greenwood [1975a] and Miller [1973]. One exception to this generalization is Niedorcorn and Bechdolt [1969] in which utility theory is used to derive a gravity model of migration.

[2]Greenwood's [1975a, p. 521] explanation for the inclusion of the change of per capita income in migration functions is that the individual may extrapolate current rates of growth of income differentials into the future. An equilibrium process determines regional wage rate relatives so extrapolations will typically not be realized. Such extrapolation also seems inconsistent with the general picture of highly imperfect information.

[3]The choice of where to go once one has decided to leave (in-migration) is somewhat more difficult to model satisfactorily and is, therefore, left to another paper. A preliminary reconnisance suggests that both changes and levels enter an in-migration function.

[4]The Southeast region's per capita personal income has risen from 52 percent of the national average in 1929 to 83 percent in 1972. The Middle Atlantic region has fallen from 138 percent to 113 percent. The coefficient of variation across states has fallen from 38.1 percent in 1940 to 20.4 percent in 1960 and 14.9 percent in 1972. [Bretzfelder 1973, p. 41]. The within state coefficient of variation for state economic areas ranges between 2 and 20 percent.

[5] These r_j's are based on a 5 percent real discount rate and the yearly probabilities of separation from the work force due to retirement and death estimated by the B.L.S. [Fullerton, 1971]. The real value of wage differentials and living quality differences is assumed constant. Assumptions that the real value of wage or living quality differences is assumed constant. Assumptions that the real value of wage or living quality differentials will grow with real per capita income or that there is a risk of their diminishing over time are easily incorporated by adjusting the discount rate up or down. If the discount rate were 10 percent 30-45 year old men would have an r_j of .955 and a dr_j of -.038. For the average man 50-65, $r_j = .69$ **and** $dr_j = -.10$.

[6] The implied residual variance is $\Sigma\left[V(L_i)\cdot\frac{1}{V(L_i)}\right] / \Sigma\frac{1}{V(L_i)}$

$$= \Sigma\frac{50}{V(L_i)} = 50/\ \Sigma\frac{1}{\frac{4}{95}+\frac{1}{N\hat{p}\ (1-\hat{p})}}.$$

$\hat{P}$ is the anti logit of the states' predicted logit from the first stage. An iterative approach to determining the relative importance of equation error is possible here but does not seem to be required given the success of the first guess.

[7] If the levels model were the true specification and **five-year** averages were the way to measure the true level, using beginning of period levels instead may be interpreted as using an imperfectly measured regressor. Auxiliary regressions were run to examine the size of the resulting errors in variables bias, if an average levels model were the true model. The estimated coefficients from model 1 would be consistent estimators of .96 of the true wage coefficient and .73 of the true unemployment coefficient. These biases are quite small and will have only marginal impacts on the significance of the variable.

[8] These variables are intended as proxies for the diversity of job opportunities and living conditions within the state and thus serve as controls for the likelihood of migrating between labor markets without crossing a state boundary. Miller [1973] has pointed out the need for such control variables. The variables used in this paper are the log of 1965 employment in the state and the weighted average of the logged population of a state's SMSAs (non SMSA portions of the state are averaged in as if they all had a population of 20,000).

[9] Since the N and $\hat{P}$ are known, it is possible to calculate a direct estimate of the variance of the measurement error of each weighting scheme. Measurement errors share of the residual variance is approximately three-fourths for operatives and one-half for craftsmen. An estimate of the equation error may be derived by subtraction. At the mean out-migration rate of .05, a one standard (equation) error shift of the logit of outmigration implies only a .6 percent change in the size of the operative labor force in 1970 and a 1.2 percent change in the size of the craftsmen labor force. Scaled in probability terms (i.e., proportionate effects on labor force growth), the variance of our equation error is approximately .00004 for operators and .00014 for craftsmen. The overall variance of employment growth rates (.00284) is more than 20 times greater.

[10] Employment growth rates were not in our original runs because they did not fit neatly into the theory and because there was a potential positive feedback simultaneity that might bias coefficients away from zero (thus making our hypotheses tests too liberal). In response to criticism of early drafts both lagged and contemporaneous growth rates were added to the model and, unexpectedly, positive coefficients were consistently

obtained. An after the fact explanation of this result is that localities with high rates of employment growth contain a high proportion of recent in-migrants. While being born in another state has been controlled, the recency of one's arrival has not, so growth rates may, in fact, be measuring the strength of local ties.

[11] These two areas of economics seem to be developing quite independently In Richardons [1973] 300 + item bibliography only four articles or books relating to migration are cited. The Borts and Stem [1964] book and the Chicken or Egg controversy started by Muth [1971] are notable expectations to this generalization.

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