The Migration Decision: What Role Does Job Mobility Play?

By ANN P. BARTEL*

An important characteristic of the *U.S.* population is its geographic mobility. In 1970, 18 percent of the population was living in a county that was different from their 1965 county of residence; half of these migrants had also moved across state lines.[1] Previous work on geographic mobility can be classified into two categories.[2] The first is composed of studies that have used aggregate data (for example, Samuel Bowles, Michael Greenwood, 1969, Ira Lowry, and Aba Schwartz) to examine the determinants of net or gross migration for *SMSAs* or other geographic divisions. The second category of research has used data on individuals (for example, Julie DaVanzo, Richard Kaluzny, John Lansing and Eva Mueller, and Solomon Polachek and Francis Horvath) to explore the relationship between an individual's characteristics and his decision to migrate.

This article continues the work on the analysis of the individual's decision to migrate, but differs from the previous studies by focusing on the relationship between job mobility and migration. First, the proportion of geographic mobility that occurs in conjunction with a job change is calculated. Second, it is shown that the true effects of human capital variables, job characteristics, and family variables on the decision to migrate are best measured when one takes account of the relationship between migration and job mobility. Third, the effect of migration on the wage gains of individuals is studied and again the need for distinguishing among moves that were associated with quits, layoffs, and transfers is clearly shown. Finally, by using three data sets that encompass different age groups (the National Longitudinal Surveys (*NLS*) of Young and Mature Men and the Coleman-Rossi Retrospective Life History Study), the importance of the relationship between migration and job mobility is demonstrated at different points in the life cycle.[3]

Section I of the article presents some summary statistics on the extent of geographic mobility among the individuals in the samples and documents the relationship between migration and job mobility. In Section II a framework for analyzing the decision to migrate is discussed. Sections III and IV present the empirical results while Section V summarizes the analysis.

I. Some Evidence on Migration and Job Mobility

Table 1 contains summary statistics on the rate of migration and job mobility in the three data sets. In the case of the two *NLS* samples migration is defined as a move to a different Standard Metropolitan Statistical Area (*SMSA*) or county, while for the Coleman-Rossi individuals migration is a move across state lines. The Coleman-Rossi migration rates are therefore expected to be low relative to intercounty migration rates. Column 1 presents data for the *NLS* Young Men for the period 1971–73. These men were between the

*Columbia University Graduate School of Business and National Bureau of Economic Research. I would like to thank Jacob Mincer and Robert Pollak for their very helpful comments on a previous draft of this paper. This research was supported through the National Bureau of Economic Research by a grant from the National Science Foundation (Grant No. SOC71-03783 A03). The opinions expressed herein are my own and do not necessarily reflect the views of the National Science Foundation.

[1]See the 1970 Census of Population, General Social and Economic Characteristics, Table 87.

[2]See Michael Greenwood (1975) for a summary of the literature on geographic mobility.

[3]The National Longitudinal Survey (*NLS*) of Mature Men is described in U.S. Department of Labor (1970–74), the *NLS* of Young Men is described in U.S. Department of Labor (1970–77), while the Coleman-Rossi data, which were collected in January 1969, are discussed in Z. Blum, N. Karweit, and A. Sorensen. The analysis is restricted to the white men in all three samples.

TABLE 1—DESCRIPTION OF GEOGRAPHIC MOBILITY

	NLS Young Men 1971–73	Coleman-Rossi 1964–69	*NLS* Mature Men 1966–71
Proportion who moved	.20	.13	.06
Proportion of moves involving quits	.49	.53	.38
Proportion of moves involving layoffs	.13	.14	.26
Proportion of moves that did not involve a job change	.38	.33	.36
Proportion of quitters who migrated	.32	.20	.16
Proportion of those laid off who migrated	.26	.19	.07
Proportion of job stayers who migrated	.12	.06	.03
Proportion of job changer migrants who moved for economic reasons	.76[a]	.75[a]	.52
Proportion of migrants who moved for economic reasons	.85[b]	.83[b]	.69[b]
Proportion of moves "caused by" the decision to change jobs	.47[c]	.50[c]	.33[c]
Sample Size	1,608	580	1,790

[a]Based on data in Lansing and Mueller for similar age groups.
[b]Includes transfers.
[c]See text for method used to calculate this statistic.

ages of 19 and 29 in 1971. In column 2 data are shown for the 1964–69 period for the Coleman-Rossi individuals who were between the ages of 26 and 35 at the start of the period. Column 3 contains data for the *NLS* Mature Men for the period 1966–71. These men were between the ages of 45 and 59 at the start of the period. A shorter time interval was used for the *NLS* Young Men in order to maximize the number of individuals who were not enrolled in school.[4]

The second, third, and fourth rows in Table 1 document the relationship between migration and job mobility. For all samples we observe that roughly two-thirds of all moves involved a job separation.[5] To determine whether these moves were in fact caused by the decision to change jobs, information on the reasons for migration is necessary. The *NLS* Mature Men data set provides information on whether a move was undertaken for economic reasons or personal (for example, family, health) reasons. One can argue that moves that involved job separations and that were made for economic reasons were in fact caused by the decision to change jobs. Those moves that involved job separations but were made for personal reasons can be said to have caused the accompanying job separations. For the *NLS* Mature Men, 52 percent of those individuals who migrated and separated indicated that they moved because of economic reasons. In order to estimate what percentage of *all* moves are *caused* by the decision to *change* jobs, however, the number of individuals who separated *and* who migrated for economic reasons as a percentage of *all* migrants is calculated. For the mature men, this proportion is one-third. For the younger samples, no information is provided on the reasons for migration. Data from another source, however, enables us to make similar calculations for these age groups.[6] As Table 1 indicates, 75 percent of those individuals who migrated and separated moved because of economic reasons resulting in one-half of all moves in the younger cohorts being *caused* by the decision to change jobs. This analysis, therefore, indicates the importance of studying the decision to migrate in conjunction with the decision to separate from a firm.

[4]For example, in 1967, 50 percent of the sample was enrolled in school, but by 1971, only 17 percent of the sample was enrolled. At the time this study was done, data for the *NLS* Young Men were available only up to 1973.

[5]Note that for the two younger samples, 80 percent of these separation-related moves are due to quits, while for the older sample 60 percent are due to quits. These differences across samples are, of course, related to the decline in the ratio of quit rates to layoff rates with age.

[6]Lansing and Mueller find that 77 percent of individuals aged 18–24 who migrated and separated during a five-year period moved because of economic reasons while the same statistic is 75 percent for men aged 25–34. See their Table 9.

II. Theoretical Framework and Empirical Specifications

The data presented in Section I document the relationship between migration and job mobility. In this section the framework within which the decision to migrate can be analyzed is discussed and it is shown how job separations can be integrated into this analysis.

Economic theory predicts that an individual will attempt to sell his services in the market which offers him the highest return. Larry Sjaastad utilized this basic concept in his analysis of internal migration in the United States. The individual is guided by his discounted net return from migrating at time t; if this net return is positive, he will migrate. In other words,

$$PM_t = f(G_t) \tag{1}$$

where PM_t is the probability that the individual moves in time period t and G_t is the discounted net gain from moving. Thus G_t can be written as follows:

$$G_t = Y_t^* - Y_t - C_t \tag{2}$$

where Y_t^* is the present value of the expected real income stream if the individual migrates in time period t, Y_t is the present value of the expected real income stream in the current location calculated at time t, and C_t are the costs of migration as well as such costs as the loss of the wife's earnings (assuming such a loss occurs), the costs of uprooting school-age children as well as the time costs of searching for a job and residence in the new location. If $G_t > 0$, the individual is assumed to migrate.

The probability PM_t can be viewed as the unconditional probability of migration. As was shown in Section I, some migrants are also job quitters, some were laid off, and the remainder are individuals who did not change employers. In other words, PM can be viewed as the sum of three joint probabilities:

$$PM = P(Q \cap M) + P(L \cap M) + P(NS \cap M) \tag{3}$$

where

$P(Q \cap M)$ = the joint probability of quitting and migrating

$P(L \cap M)$ = the joint probability of being laid off and migrating

$P(NS \cap M)$ = the joint probability of not separating from the firm and migrating

Moreover each joint probability can be rewritten as

$$P(X \cap M) = P(X) \cdot P(M \mid X) \tag{4}$$

where X = Q, L, or NS,[7] $P(X)$ is the probability of a quit, layoff, or no separation, and $P(M \mid X)$ is the probability of migration conditional on a quit, layoff, or no separation. Equation (4) shows how the decision to migrate is directly linked to the probability of a job separation. In studying migration, therefore, we see that it is crucial to have an understanding of both the process of job mobility and the determinants of the conditional probabilities of migration.

In order to study the determinants of the probability of migrating, those variables which measure the discounted net return from moving must be identified. Since the decision to move has already been shown to be closely tied to the decision to change jobs, the analysis is in part an attempt to measure the discounted net return from changing jobs. In addition, those variables which determine the conditional probability of migration must be examined. Of course, there may be some overlap in the sets of variables that determine $P(X)$ and $P(M \mid X)$. More important is the fact that some variables may affect the discounted net return from migration only because they affect the discounted net return from separating. This points out the importance of examining the joint probabilities of migration rather than the unconditional probability of migrating. Moreover, it suggests that a convenient way of determining whether

[7]A move that did not involve a job separation could either be an intrafirm job change or residential mobility. Although the data do not distinguish between these two types of moves, the fact that migration is defined as a move to a different *SMSA*, county, or state indicates that most of these moves are probably transfers. In the remainder of the article, this assumption is maintained.

a variable's measured effect on migration is due solely to its effect on the probability of changing jobs is to compare the effect of that variable on the probability of separating and migrating with its effect on the probability of separating and not migrating (i.e., changing jobs in the local labor market).

For example, consider the effect of the individual's current wage. According to the theory of specific training, the wage should have a negative effect on the probability of quitting, but an ambiguous effect on the probability of a layoff.[8] If the wage does not affect the conditional probability of migration, then a negative effect of the wage on the probability of migrating may be observed only if the individual quit his job. And, if the wage affects migration simply because it affects the probability of separating, then the measured effect of the wage on the joint probability of migrating and separating would be the same as its effect on the probability of separating in the local labor market. In the case of transfers, however, the wage should have a positive effect on migration since employers would be likely to transfer (i.e., invest in) those individuals who already have a large amount of specific training and are closely tied to the firm.[9] This analysis shows, therefore, that the true effect of the wage on the probability of migration can only be estimated when the different types of moves are distinguished from one another.

Other variables can be suggested as determinants of the discounted net gain from moving. Education should have a positive effect on the conditional probability of migration since more highly educated individuals would tend to have better information about nonlocal job opportunities, may be more adaptable to change, and tend to be in occupations that operate in a national labor market. This would predict a stronger effect of education on nonlocal separations than local separations. Whether education will in fact have an effect on the unconditional probability of migration, however, is unclear. For example, Borjas and I present evidence that more educated individuals are significantly less likely to be laid off. This would suggest that in the case of the joint probability of being laid off and migrating, the effect of education is ambiguous in sign. Therefore in estimating the relationship between education and migration one would want to distinguish among types of moves.

One of the most important sets of determinants of the net return from migration is the characteristics of the individual's family.[10] For example, married individuals with working wives should have higher costs of migration than those whose wives are not in the labor force. Similarly, individuals with school-age children should have a lower net return from migration, everything else held constant. Again, however, the effects of these variables on the decision to migrate may depend on the association of a move with a job separation. For example, the presence of a working wife may have little effect on the probability of being transferred by one's employer; the true inhibiting effect of wife's

[8] The theory argues that employees with more worker-financed specific training are less likely to quit and those with more firm-financed specific training are less likely to be laid off. Following Donald Parsons, since an individual's wage can be expressed as

$$W = a_0 + a_1E + a_2S_w$$

where E = education and S_w = worker-financed specific training, the quit probability will be inversely related to the wage when education is held constant. The sign of the relationship between the layoff probability and the wage depends on whether firm-financed specific training is positively or negatively correlated with worker-financed specific training. Although the positive correlation is more likely, the layoff probability may still be positively related to the wage if job instability is compensated by a wage premium. See Robert Hall for a discussion of the relationship between wages and separation rates according to the theory of compensating wage differentials. George Borjas and I discuss additional theories that can be used to explain the wage rate-separation rate relationship.

[9] The equation in fn. 8 shows that the individual's wage is positively correlated with the amount of specific training he possesses since an individual with more worker-financed specific training is also likely to have more specific training in total. In other words, the incentives that exist for the worker to invest in specific training are also likely to induce the firm to invest in the worker.

[10] See DaVanzo, Larry Long, Jacob Mincer, and Polachek and Horvath for empirical evidence.

labor force participation on the migration decision of job separators would then not be correctly estimated in an analysis that did not distinguish among types of moves.

This approach also indicates why certain job-related characteristics should affect the net return from migration. For example, individuals with low levels of tenure in the current job are more likely to experience a job separation.[11] And, given the underlying relationship between migration and job separations, these individuals would therefore be likely to move geographically. More important is the notion that the correlation between job tenure and length of residence may produce the observed negative effect of residence when in fact the true causative variable is job tenure.

The analysis presented here thus shows that the relationship between job mobility and migration can be demonstrated by estimating the following set of linear probability equations:

$$PM_i = a(Z_i, F_i, J_i) \tag{5}$$

$$\begin{aligned} P(Q \cap M)_i &= b(Z_i, F_i, J_i) \\ P(L \cap M)_i &= c(Z_i, F_i, J_i) \\ P(NS \cap M)_i &= d(Z_i, F_i, J_i) \end{aligned} \tag{6}$$

$$\begin{aligned} P(Q \cap NM)_i &= e(Z_i, F_i, J_i) \\ P(L \cap NM)_i &= f(Z_i, F_i, J_i) \end{aligned} \tag{7}$$

where Z_i is a vector of individual characteristics, F_i is a vector of family characteristics, J_i is a vector of job-related characteristics, and NM means not migrating. It has been argued that the coefficients in equation (5) will not correctly estimate the effects of the independent variables for all movers since, as equations (6) show, there are three distinctly different types of moves. Further, a comparison of equations (6) and (7) will show whether an independent variable affects the probability of migrating simply because it determines the probability of separating, that is, whether it is useful to distinguish local separations from nonlocal separations.

III. Empirical Findings on the Determinants of Migration

In this section the results of estimating equations (5), (6), and (7) using data from the *NLS* of Young Men, the Coleman-Rossi Retrospective Life Histories Study, and the *NLS* of Mature Men are presented. Table 2 contains the regressions for the *NLS* Young Men sample while Table 3 contains similar regressions for the Coleman-Rossi sample and Table 4 presents the regressions for the *NLS* Mature Men.[12,13] The regressions in these tables do not hold job tenure and length of residence constant since it can be argued that these variables are serially correlated with the dependent variable, that is, previous moves have determined current job tenure and length of residence. In fact, when tenure and residence are added to the regressions, some of the other independent variables do become less significant (but the conclusions of this analysis are unchanged) indicating that these variables also determined previous mobility. The coefficients on tenure and residence from these regressions are shown in Table 5 and the complete regressions are available from me upon request. In order to focus on the distinction between unconditional and joint probabilities of migration, each independent variable is discussed in turn to show how its measured effect on migration depends on the associated job separation. The variables are defined in Table 2; note that the independent variables are measured at the beginning of the period under study.

A. *The Wage*

As discussed in Section II, the effect of the individual's wage rate on the probability of migrating may depend on whether migration

[11] See the papers by Borjas and myself and by Boyan Jovanovic and Mincer.

[12] For the *NLS* Young Men the time period under study is 1971–73, for the Coleman-Rossi sample it is 1964–69, and for the *NLS* Mature Men it is 1966–71. See fn. 4 for a discussion of the reason that the *NLS* Young Men analysis was restricted to a two-year period.

[13] Since the dependent variables are dichotomous, ordinary least squares is not the proper estimating technique. This article therefore utilizes maximum likelihood logit; the coefficients presented in the tables are the marginal

TABLE 2—DETERMINANTS OF 1971–73 MIGRATION: *NLS* YOUNG MEN[a,b]
MAXIMUM LIKELIHOOD LOGIT ESTIMATES

	Uncond	*Migqt*	*Miglay*	*Migtr*	*Nmquit*	*Nmlay*
EDUC	.0127	.0057	.0023	.0044	−.0222	−.0064
	(2.37)	(1.45)	(1.08)	(1.25)	(−3.98)	(−1.79)
EXPER	−.0126	−.0102	−.0018	−.0011	−.0070	−.0003
	(−3.18)	(−3.36)	(−1.09)	(−.45)	(−1.84)	(−.11)
WAGE	.0060	−.0059	.0006	.0099	−.0136	−.0049
	(.89)	(−1.10)	(.22)	(2.43)	(−1.73)	(−.97)
MAR	−.0223	.0155	−.0088	−.0265	−.0381	−.0263
	(−.80)	(.76)	(−.79)	(−1.42)	(−1.34)	(−1.47)
WLFP	.0526	.0035	.0043	.0419	.0438	−.0248
	(1.68)	(.15)	(.32)	(2.09)	(1.35)	(−1.09)
WINC	−.0009	−.0002	−.0001	−.0006	−.0002	.0003
	(−1.51)	(−.45)	(−.41)	(−1.46)	(−.36)	(.67)
SCHL	.0071	.0185	−.0138	−.0029	−.0392	.0127
	(.23)	(.78)	(−.84)	(−.15)	(−1.18)	(.61)
HLTH	.0692	.0024	−.0013	.0568	−.0078	.0090
	(1.98)	(.09)	(−.08)	(2.87)	(−.19)	(.36)
UNEMP	−.0048	.0181	.0161	−.0529	.0659	.0587
	(−.20)	(1.09)	(1.91)	(−2.82)	(2.97)	(4.27)
χ^2	50.50	32.41	16.19	34.57	63.78	42.37
N	1608	1608	1608	1608	1608	1608

[a]Asymptotic t-ratios are given in parentheses. Definition of variables are *EDUC* = years of education, *EXPER* = potential experience (as of 1971) since completion of schooling (*NLS* Young Men), *REM* = time remaining until retirement as of 1966 (*NLS* Mature Men), *AGE* = age in 1964 (Coleman-Rossi), *WAGE* = hourly (*NLS*) or monthly (Coleman-Rossi) wage at the beginning of the period, monthly wage is in tens of dollars, *MAR* = one if individual is married, *WLFP* = one if individual's wife was in the labor force at the beginning of the period under study, *WW* = wife's hourly wage (*NLS* Mature Men), *WINC* = wife's earnings in hundreds of dollars (*NLS* Young Men), *SCHL* = one if individual has school-age children, *HLTH* = one if individual's health limits kind or amount of work (*NLS* Young and Mature Men), *UNEMP* = one if individual unemployed during the previous year, *JOB* = job tenure in years at the beginning of the period and *RTEN* = the difference between length of residence and job tenure at the beginning of the period.

[b]Column headings refer to the probability of migration: *Uncond* is the unconditional probability; *Migqt* is the joint probability of migrating and quitting; *Miglay* is the joint probability of migrating and being laid off; *Migtr* is the joint probability of migrating and not changing employers; *Nmquit* is the joint probability of not migrating and quitting; *Nmlay* is the joint probability of not migrating and being laid off.

is associated with a job separation. The results in the first column of each table show that the wage has no effect on the unconditional probability of migration. The reason for this somewhat paradoxical result is made clear by an examination of the other regressions in Tables 2, 3, and 4. We find that *WAGE* has a negative effect (which is significant only for the *NLS* Mature Men) on the probability of migrating and quitting in all three samples. However, in the case of the joint probabilities of migrating and being laid off or migrating and not changing employers, the wage coefficient is always nonnegative and in some cases is significant. The reason for this nonnegative wage effect was suggested in Section II. Since the joint probability of migrating is a function of the probability of separating, the nonnegative wage coefficient may be due to a nonnegative wage effect on the probability of being laid off and the probability of being "promoted" via a transfer. In fact, for the younger cohorts (both *NLS* and Coleman-Rossi), transfers depend positively and significantly (in the case of the *NLS* Young Men) on the wage level.

effects of the independent variables on the dependent variable, evaluated at the mean of the dependent variable.

Is the negative effect of the wage on the probability of migrating and quitting due solely to the negative relationship between

TABLE 3—DETERMINANTS OF 1964–69 MIGRATION: COLEMAN ROSSI[a] MAXIMUM LIKELIHOOD LOGIT ESTIMATES

	Uncond	*Migqt*	*Miglay*	*Migtr*	*Nmquit*	*Nmlay*
EDUC	.0136	.0049	.0008	.0082	−.0048	−.0073
	(2.62)	(1.25)	(.41)	(2.70)	(−.74)	(−2.07)
AGE	−.0009	−.0010	−.0002	.0002	−.0014	.0003
	(−2.04)	(−2.76)	(−1.09)	(.90)	(−2.33)	(.85)
WAGE	.0003	−.0006	.0001	.0004	−.0006	.0002
	(.46)	(−1.08)	(.79)	(1.52)	(−.71)	(.53)
MAR	−.0371	−.0021	−.0096	−.0222	−.0381	.0045
	(−.96)	(−.07)	(−.71)	(−.94)	(−.67)	(.13)
WLFP	−.0210	−.0253	−.0065	.0119	.0749	−.0167
	(−.58)	(−.92)	(−.43)	(.57)	(1.68)	(−.59)
SCHL	.0010	.0083	−.0078	.0026	.0404	−.0043
	(.03)	(.31)	(−.50)	(.13)	(.90)	(−.17)
UNEMP	.1808	.0646	.0452	.0294	−.1624	.0986
	(2.73)	(1.26)	(2.99)	(.68)	(−1.04)	(2.08)
χ^2	21.24	13.45	10.15	14.81	12.97	8.72
N	579	579	579	579	579	579

[a]Asymptotic *t*-ratios are given in parentheses. Variables are defined in Table 2.

wages and quitting? This question can be answered by looking at the regressions on the probability of quitting and not migrating. The results show that the wage effect in these equations is as strong or stronger than the effect in the associated migration equations. The stronger effect in the local quit equations is due in part to the larger mean value for local quits[14] which is then applied to the logit coefficients to estimate marginal effects (see

[14]The mean value for *NMQUIT* is at least twice as large as the mean value for *MIGQT* in all three samples.

TABLE 4—DETERMINANTS OF 1966–71 MIGRATION: *NLS* OLDER MEN[a] MAXIMUM LIKELIHOOD LOGIT ESTIMATES

	Uncond	*Migqt*	*Miglay*	*Migtr*	*Nmquit*	*Nmlay*
EDUC	.0072	.0038	−.0010	.0052	.0029	−.0064
	(3.85)	(3.24)	(−1.02)	(4.58)	(1.15)	(−2.38)
REM	.0017	.0010	.0014	−.0007	.0005	.00007
	(1.33)	(1.29)	(2.21)	(−.91)	(.32)	(.04)
WAGE	−.0032	−.0114	.0015	−.0003	−.0229	.0030
	(−1.01)	(−3.40)	(1.49)	(−.19)	(−3.70)	(.73)
MAR	−.0131	−.0074	−.0056	.0198	−.0326	−.0034
	(−.64)	(−.62)	(−.60)	(.95)	(−1.09)	(−.10)
SCHL	−.0213	−.0090	−.0110	−.0013	.0092	−.0265
	(−1.71)	(−1.09)	(−1.63)	(−.19)	(.56)	(−1.52)
WLFP	−.0192	−.0065	−.0118	−.0017	.0385	−.0086
	(−1.34)	(−.75)	(−1.35)	(−.17)	(1.89)	(−.43)
WW	.0012	.0029	.0001	−.0019	−.0006	.0026
	(.38)	(1.89)	(.04)	(−.53)	(−.10)	(.51)
HLTH	−.0110	.0013	−.0083	.0016	−.0239	.0401
	(−.79)	(.15)	(−1.25)	(.17)	(−1.31)	(1.81)
UNEMP	.0083	−.0045	.0128	−.0141	.0934	.1454
	(.40)	(−.32)	(1.61)	(−.70)	(4.44)	(6.84)
χ^2	23.51	25.26	16.58	29.52	50.51	58.68
N	1790	1790	1790	1790	1790	1790

[a]Asymptotic *t*-ratios are given in parentheses. Variables are defined in Table 2.

fn. 13). It can therefore be concluded that the relationship between wages and migration is strongly dependent on the fact that job separations accompany migration; the only case in which a move is seen to be negatively related to the wage (i.e., quitting and migrating) is found to be due entirely to the negative effect of the wage on the job separation itself.

B. *Education*

Education has a positive and significant effect in all samples on the unconditional probability of migrating. This result is consistent with the findings of other studies surveyed in Greenwood (1975) and has been explained as being due to the more educated individual's ability to adapt to new locations and his greater efficiency in searching for jobs in other locations. The empirical results in this paper show that although education is not positive and significant in all of the joint-probability migration equations, it does have an effect on migration that is independent of its effect on the probability of separating. This can be seen by comparing the coefficients in the local and nonlocal separation equations; in all cases *EDUC* is more positive in the case of a nonlocal separation. Unlike the wage, education does have an independent effect on the decision to migrate.

C. *Family Variables*

The costs of migration that are usually associated with marital status can be measured by information on the wife's labor force participation and the ages of the children. The effect of the wife's labor force status is measured by a dummy variable indicating her participation in the labor force (*WLFP*) and a continuous variable measuring her wage or annual earnings (*WW*, *WINC*).[15] For men in their 30's, 40's, or 50's (Tables 3 and 4) we find that wife's labor force participation has a negative but insignificant effect on the unconditional probability of migration.[16] This occurs for two reasons. First, wife's participation has no effect in these samples on the probability of being transferred. Second, although wife's participation does inhibit migration in the case of job quitters, this effect can not be directly observed in the joint-probability (*MIGQT*) equations. The reason is that wife's participation has a positive and significant effect on the probability of quitting locally. Therefore, to measure the negative effect of *WLFP* on the probability of migrating, one should compare the coefficients in the *MIGQT* and *NMQUIT* equations. As can be seen from Tables 3 and 4, *WLFP* is less positive in the *MIGQT* equations, reflecting the inhibiting effect of this variable on quitting into another labor market.[17] Similarly, the negative effect of school-age children (*SCHL*) on the migration decision of job quitters in Table 4 is shown by comparing the effects of this variable on *MIGQT* and *NMQUIT*. This analysis therefore shows the importance of decomposing the unconditional probability of migration in order to correctly estimate the extent to which a working wife and the presence of school-age children inhibit a job quitter from changing locations.

D. *Job Tenure and Length of Residence*

Previous research on migration has found that one of the most important determinants of the decision to migrate is the length of

[15]In the case of the Coleman-Rossi sample, many men whose wives worked failed to report their wages or earnings. Since there were so many missing values for this variable, it was deleted from the regressions for this sample.

[16]In the case of men who are in their 20's (Table 2), the effect of wife's labor force participation on the unconditional probability of migration depends on the amount of the wife's earnings. If the wife's annual earnings are below $5,800 her participation does not inhibit migration; at earnings levels above $5,800 her participation has a negative effect which eventually becomes significant. (This is calculated by realizing that *WINC* is actually an interaction term between the dummy variable *WLFP* and the wife's earnings if she works.) Since these women are in their childbearing years and are likely to participate intermittently in the labor force, their current participation is not an inhibiting factor in migration unless their earnings represent a substantial contribution to family income.

[17]This result also holds for the young men in Table 2.

residence in the current location.[18] Individuals who have lived in the current location a long time may be less likely to move because they have built up a stock of capital that is specific to this location; that is, over time, strong community ties will have been developed thereby raising the costs of migration. One must also recognize, however, that the negative effect of length of residence may be due to the relationship between residence and job tenure. To the extent that the individual has not changed jobs during his stay in this location, job tenure and length of residence will be strongly correlated. Moreover, there is substantial evidence that job tenure reduces the probability of a job separation because of the positive correlation between tenure and job-specific training.[19] Since the relationship between migration and job separation has already been documented, it is quite possible that the observed negative effect of length of residence on migration may be due to the negative effect of tenure on job separations. We would like to be able to identify whether there is an independent effect of residence on migration.

Fortunately, since the data sets provide information on both length of residence and job tenure, the separate effects of length of residence and job tenure on the decision to migrate can be identified. This is accomplished by defining a variable *RTEN* which equals the difference between length of residence and job tenure and including this variable as well as *JOB* (length of job tenure) in the regressions. Then *RTEN* captures the effect of a year of residence net of job tenure, that is, the "pure" residence effect, while the coefficient on *JOB* is the sum of the pure residence effect and the pure job effect, if it exists. If *RTEN* has a significant effect on the decision to migrate and the coefficient on *JOB* exceeds (in absolute value) the coefficient on *RTEN*, then it can be concluded that the inhibiting effect of residence observed in other studies is due to the acquisition of *both* location-specific capital and job-specific capital.

Table 5 contains the estimated coefficients on *JOB* and *RTEN* for each of the unconditional and joint-probability equations. The results show that *RTEN* has a significant negative effect in the equations referring to the young cohorts (*NLS* Young Men and Coleman-Rossi); thus for these samples length of residence has an inhibiting effect on migration which is net of the relationship between residence and tenure. For the older men, however, the correlation between residence and job tenure is very high; *RTEN* only ranges from zero to nine years. For these men, *RTEN* is insignificant indicating that the negative relationship between residence and migration is due solely to the negative effect of job tenure on separations. It is important to note, however, that for all three samples,

TABLE 5—MAXIMUM LIKELIHOOD LOGIT COEFFICIENTS AND ASYMPTOTIC *t*-RATIOS ON *RTEN* AND *JOB* FOR ALL SAMPLES[a]

Dependent Variable	*RTEN*	*JOB*
NLS Young Men		
Unconditional	−.0088	−.0239
	(−8.16)	(−4.31)
Migrate and quit	−.0044	−.0244
	(−5.47)	(−5.01)
Migrate and be laid off	−.0018	−.0154
	(−3.81)	(−3.92)
Transfer	−.0027	.0029
	(−4.05)	(1.00)
Coleman-Rossi		
Unconditional	−.0240	−.0204
	(−5.76)	(−3.75)
Migrate and quit	−.0086	−.0100
	(−2.94)	(−2.42)
Migrate and be laid off	−.0022	−.0054
	(−1.43)	(−1.79)
Transfer	−.0156	−.0072
	(−2.21)	(−2.21)
NLS Mature Men		
Unconditional	−.0016	−.0042
	(−.80)	(−5.86)
Migrate and quit	−.0006	−.0042
	(−.45)	(−5.03)
Migrate and be laid off	−.0002	−.0013
	(−.15)	(−3.09)
Transfer	−.0011	−.0002
	(−.85)	(−.73)

[a]*t*-ratios shown in parentheses.

[18]For example, see Kaluzny and Polachek and Horvath. Recall that this observed effect may be due in part to serial correlation in the dependent variable.

[19]See the papers by Borjas and myself and by Jovanovic and Mincer.

TABLE 6—COEFFICIENTS ON MIGRATION DUMMY VARIABLES FROM WAGE GROWTH REGRESSIONS[a]

	NLS Young Men 1971–73			Coleman-Rossi 1964–69			*NLS* Mature Men 1966–71		
GEOG	.3621			−.9982			−.0095		
	(2.32)			(−.03)			(−.06)		
MIGQT		.4820	.5341		−39.76	−53.14		−.0849	−.1502
		(2.27)	(2.44)		(−1.16)	(−1.50)		(−.34)	(−.59)
MIGLAY		−.3147	−.2572		−129.80	−143.10		−.1628	−.2233
		(−.82)	(−.67)		(−1.97)	(−2.16)		(−.57)	(−.77)
MIGTR		.4338	.4792		119.17	106.92		.1683	.1327
		(1.89)	(2.05)		(2.73)	(2.42)		(.69)	(.54)
NMQUIT			.1379			−36.14			−.2838
			(.85)			(−1.80)			(−2.17)
NMLAY			.1514			−4.39			−.1059
			(.63)			(−.41)			(−.83)

[a]Definitions of variables are *GEOG* equals one if individual migrated during the period; *MIGQT* equals one if individual quit and migrated; *MIGLAY* equals one if individual was laid off and migrated; *MIGTR* equals one if individual migrated but did not change employers; *NMQUIT* equals one if individual quit but did not migrate and *NMLAY* equals one if individual was laid off but did not migrate.

when a job separation accompanies a geographic move, job tenure *itself* reduces the probability of migration, that is, the coefficient on *JOB* exceeds that on *RTEN*, because of the relationship between tenure and job separations. In the case of a transfer, however, the pure job effect (coefficient on *JOB* minus coefficient on *RTEN*) is actually positive; employers appear to be more likely to "promote" those individuals who have shown a commitment to the firm. This analysis therefore shows that the negative effect of residence on migration observed in other studies is misleading in two respects. First, when a job separation accompanies a geographic move, part of the inhibiting effect of residence is due to the negative effect of tenure on separations. Second, when the geographic move is an intrafirm transfer, the effect of residence may be nonnegative if tenure is not held constant since job tenure increases the probability of a transfer.

IV. Wage Gains from Migration

Previous work on migration has not conclusively established that migrants have larger wage gains than individuals of similar characteristics who do not migrate.[20] Since this article has already shown that it is important to distinguish among types of moves in examining the determinants of migration, the distinction may also help in obtaining a more accurate measure of the return to migration. Table 6 contains coefficients on dummy variables measuring migrant status from regressions on absolute wage growth for each of the three samples.[21] The migrant status dummy variables are defined in the footnote to the table. While a vector of standardizing variables was included in the wage growth regressions, these coefficients are not reported here.[22]

[20]Greenwood (1975) shows that while many studies have found a positive return to migration, others have been unable to support this conclusion. He argues that the return to migration can be correctly calculated only if the migrant population is disaggregated as finely as possible; in other words, the return to migration differs appreciably across groups. This paper suggests that job mobility may be an important characteristic by which migrants should be stratified.

[21]By analyzing the effects of migration on wage growth rather than wage levels, we avoid the possibility for simultaneity bias in the wage equation. If there are certain unobserved personal characteristics which affect both an individual's wage and his decision to migrate, a wage level equation will be biased. A wage growth equation nets out these unobserved individual differences which affect an individual's earnings throughout the life cycle.

[22]The vector includes education, years of experience, marital status, wife's labor force status and income, presence of school children, job tenure, length of residence, and unemployment experience.

The results in Table 6 show that if no distinction is made among types of moves, a positive and significant effect of migration on wage growth is observed only for the *NLS* Young Men. Distinguishing among moves related to quits, moves related to layoffs, and transfers provides a more revealing picture of the returns to migration. For men in their 20's and 30's (*NLS* Young Men and Coleman-Rossi), transfers have a positive and significant effect on wage growth. In other words, young men who are transferred by their employers achieve wage gains that are substantially larger than the gains of men with similar characteristics who do not migrate. It therefore appears that in this age group a transfer acts as a promotion within the firm. While men in their 50's who are transferred do not receive wage gains that are significantly larger than that of the nonmigrants, it is important to note that this type of move results in the largest wage gain (the coefficient on *MIGTR* is larger than those on *MIGQT* and *MIGLAY*).

Of all the coefficients on the separation-related moves, only one is significant: the *NLS* Young Men who quit and migrate achieve significantly larger wage increases than nonmigrants. Does this imply that for the two older cohorts a geographic move that accompanies a job separation does not pay? The answer to this question depends on with whom the migrant is being compared. For example, in all three samples, individuals who quit and migrate do better than individuals who are laid off and migrate (compare *MIGQT* and *MIGLAY*).[23,24] Further, in the *NLS* Mature Men sample, individuals who quit and migrate achieve larger wage gains than individuals who quit but do not migrate (compare *MIGQT* and *NMQUIT*). In general, however, one can conclude that of the three types of moves, transfers result in the largest payoffs.

[23]Note that for layoffs this does not hold; individuals who are laid off do better if they do not migrate.

[24]The reader may be puzzled as to why a significant positive return to local quitting is not observed. The paper by Borjas and myself shows that for the *NLS* Mature Men, a substantial proportion of the quits in this age group either result in increased job satisfaction but not increased money wages or are caused by exogenous factors such as health or family problems. For the young men, only those individuals who said they quit because they found a better job had significantly larger wage gains in the 1971–73 time period. Individuals who quit because of personal problems or because of dissatisfaction with their current jobs did not have larger wage gains than stayers.

V. Summary and Conclusions

This article has analyzed the determinants and consequences of migration at different stages in the life cycle. The theme of the article has been that migration is closely related to job mobility (in fact, between one-third and one-half of all moves are *caused* by the decision to *change* jobs) and that when the decision to migrate or the returns to migration are explored, one must take account of this relationship. Several findings support this argument:

1) Economic theory predicts that, *ceteris paribus*, the wage should have a negative effect on the decision to migrate. This article shows that the wage has a significant negative effect only in the case of the joint probability of migrating and quitting. Moreover, this negative coefficient is entirely due to the negative effect of the wage on the job separation itself.

2) The true inhibiting effect of a working wife on a man's decision to migrate is shown to be correctly estimated only when the unconditional probability of migrating is decomposed; this occurs because of conflicting effects of this variable on the transfer decision, the decision to change jobs in the local market, and the decision to quit and migrate. Similar problems exist for measuring the effect of the presence of school-age children.

3) Previous research on migration has found that one of the most important determinants of the decision to migrate is the length of residence in the current location. This article shows that since residence and job tenure are positively correlated, the effect of residence on migration is at least partially due to the relationship between job tenure and the decision to change jobs.

4) The wage gains from migration are

also seen to depend on the nature of the move and the age of the migrant. Of the three types of moves, transfers in general lead to the largest wage gains; this effect is significant, however, only for the two younger cohorts. A quit-related move is also found to lead to larger payoffs than a layoff-related move for all three samples.

In conclusion, the empirical findings presented in this article support the initial argument that one must take account of job mobility in studying the determinants and consequences of the decision to migrate. The results indicate that there is an important link between the decision to migrate and the probability of a job separation; an analysis of migration that ignores this link may fail to understand the role played by many socioeconomic variables in the migration process.

REFERENCES

A. Bartel and G. Borjas, "Middle-Age Job Mobility: Its Determinants and Consequences," in Seymour Wolfbein, ed., *Men in the Pre-Retirement Years*, Philadelphia 1977.

Z. D. Blum, N. L. Karweit and A. B. Sorensen, "A Method for the Collection and Analysis of Retrospective Life Histories," rept. no. 48, Johns Hopkins Univ. Center Stud. Soc. Organ. Sch. 1969.

S. Bowles, "Migration as Investment: Empirical Tests of the Human Investment Approach to Geographical Mobility," *Rev. Econ. Statist.*, Nov. 1970, *52*, 356–62.

J. DaVanzo, "Why Families Move: A Model of the Geographic Mobility of Married Couples," R-1972, Rand Corp., Sept. 1976.

M. Greenwood, "An Analysis of the Determinants of Geographic Labor Mobility in the United States," *Rev. Econ. Statist.*, May 1969, *51*, 189–94.

———, "Research on Internal Migration in the United States: A Survey," *J. Econ. Lit.*, June 1975, *13*, 397–433.

R. Hall, "Why is the Unemployment Rate So High at Full Employment?," *Brookings Papers*, Washington 1970, *3*, 369–410.

B. Jovanovic and J. Mincer, "Labor Mobility and Wages," paper presented at Nat. Bur. Econ. Res. Conference on Low Income Labor Markets, Chicago, June 1978.

R. Kaluzny, "Determinants of Household Migration: A Comparative Study by Race and Poverty Level," *Rev. Econ. Statist.*, Aug. 1975, *57*, 269–74.

John B. Lansing and Eva Mueller, *The Geographic Mobility of Labor*, Ann Arbor 1967.

L. H. Long, "The Influence of Number and Ages of Children on Residential Mobility," *Demography*, Aug. 1972, *9*, 371–82.

Ira S. Lowry, *Migration and Metropolitan Growth*, San Francisco 1966.

J. Mincer, "Family Migration Decisions," *J. Polit. Econ.*, Oct. 1978, *86*, 749–73.

D. Parsons, "Specific Human Capital: An Application to Quit Rates and Layoff Rates," *J. Polit. Econ.*, Nov./Dec. 1972, *80*, 1120–43.

S. Polachek and F. Horvath, "A Life Cycle Approach to Migration: Analysis of the Perspicacious Peregrinator," in Ronald Ehrenberg, ed., *Research in Labor Economics*, Vol. 1, Greenwich 1977.

A. Schwartz, "Interpreting the Effect of Distance on Migration," *J. Polit. Econ.*, Sept./Oct. 1973, *81*, 1153–69.

L. A. Sjaastad, "The Costs and Returns of Human Migration," *J. Polit. Econ.*, Oct. 1962, *70*, 80–93.

U.S. Department of Labor, *The Pre-Retirement Years*, Vols. 1–4, Manpower Res. Monog. No. 15, Washington 1970–74.

———, *Career Thresholds*, Vols. 1–6, Manpower Res. Monog. No. 16, Washington 1970–77.