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QUITS, MOVES, SPATIAL EQUILIBRIUM AND WORKPLACE RELOCATION

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ABSTRACT

When worker commutes are suboptimal, quits and moves are related. Either a quit, a move, or both can achieve an optimal commute. However, with fixed costs to quitting and moving, a quit or move alone is more likely than both together. Payroll records of a firm which relocated from the central business district to a suburb of a major metropolitan area confirm this. They demonstrate that white employees rarely quit and move at the same time. Simultaneous bivariate probit estimates of move and quit behavior demonstrate that uncontrolled shocks to quits and moves are negatively correlated. Furthermore, during the spatial dislocation caused by the firm's relocation, quits and moves were direct substitutes. Employees who quit were approximately 29% less likely to move. Those who moved were approximately 40% less likely to quit.

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Quits and moves are discrete reactions to a variety of individual circumstances. In one circumstance, that of spatial disequilibrium, they are related. Workers whose commute is unsatisfactory have three possible remedies. They can move to find a residence nearer their workplace, they can quit to find a workplace more convenient to their residence, or they can both quit and move. This paper investigates the considerations under which workers choose between the three.

Changes in residence location change commuting costs and housing prices in opposite directions. A utility-maximizing model of the worker/consumer demonstrates that the optimal residence location balances the opposing gains and losses. Similarly, at the optimal workplace location, opposing changes in commuting costs and wages caused by changes in workplace location also balance. In the absence of fixed costs, workers would adjust both workplace and residence locations, at their margins, to changes in optimal commute distance.

However, if moving or quitting incurs fixed costs, commutes must be substantially suboptimal before workers would adjust them at all. If an adjustment is necessary, workers are more likely to accomplish it through a move or quit alone, rather than both. The payroll records of a single company, covering all white employees over seven years, demonstrate that quits and moves alone are much more frequent than quits and moves together.

Multivariate analyses confirm that quits and moves are, to some extent, spatial 'substitutes'. Simultaneous bivariate probit estimates of move and quit propensities demonstrate a pervasive negative correlation between the effects of random shocks on quit and move probabilities. Furthermore, they demonstrate that, following the firm's relocation from the central business district to a suburb of the same large metropolitan area, quits and moves substituted directly. Employees were 40% less likely to leave the company if they moved, and 29% less likely to change their residence if they quit.

I. The Theory of Residence and Workplace Choice

Within an urban area, moves -- residential relocations -- may take place to accomodate changes in family structure, social status, investment preferences or neighborhood amenities. Quits -- workplace relocations -- may enable professional advancement and occupational change, or derive from choices with regard to labor force participation and work environment. Many of the stimuli for moves neither arise from nor affect attitudes towards employer and workplace. Similarly, attitudes towards residence and residence location are frequently independent of the conditions which induce quits.

However, quit and move decisions may not be independent, for two reasons. First, they may be correlated if both depend on the same personal characteristics. Second, they depend directly upon each other because their relationship determines the conditions of spatial equilibrium. This second dependency is the subject of this paper.

If a worker/consumer is in spatial disequilibrium, quits and moves are obvious remedies. Either quits or moves, individually, are sufficient to reestablish spatial equilibrium. Without fixed costs, marginal adjustments to both residence and workplace locations adjust suboptimal commute distances. With fixed adjustment costs, ordinarily only one or the other will be chosen.

A simple model of utility maximization for employed consumers demonstrates the important considerations in the maintenance of spatial equilibrium in an urban area. Utility is a function of housing consumption, h, leisure, L, and an index of consumption for all other goods, x:

$$U = U(x,h,L) \quad . \tag{1}$$

Leisure is the time remaining from the time endowment, L_0, after work hours, e, and commuting time, c:

$$L = L_0 - e - c \quad . \tag{2}$$

Distances from the city center represent residential location, r^h, and workplace location, r^w. The city is circular, uniform along the circle at any distance from the center. Therefore, all workers choose residence and workplace along the same ray from the city center. Commuting time, c, is an increasing function of residential location and a decreasing function of workplace location:

$$c = c(r^h, r^w),$$

with $c_1 > 0$, $c_{11} < 0$, $c_2 < 0$, $c_{22} > 0$.

The index commodity, x, is also the numeraire. Out-of-pocket commuting costs are p_c per unit commuting time. Housing prices and wages vary spatially. Prices per unit of housing services fall with distance from the city center (Muth):

$$p_h = p_h(r^h),\ p_h' < 0\ ,\ p_h'' > 0\ .$$

Compensation per unit time at work also falls with distance from the city center: [1]

$$w = w(r^w),\ w' < 0,\ w'' < 0\ .$$

These prices define the monetary budget constraint:

$$p_c c + x + p_h(r^h)h = w(r^w)e\ .$$

This equation, in combination with equation 2, yields the overall budget constraint:

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[1] Muth and Straszheim offer theoretical demonstrations of this property. Straszheim, Eberts and Madden provide empirical verification.

$$w(r^w)L_0 = [w(r^w) + p_c]c(r^h, r^w) + w(r^w)L + p_h(r^h)h + x \ . \tag{3}$$

The optimizing worker/consumer maximizes utility (equation 1) subject to the budget constraint (equation 3) by forming a Lagrangian expression with the two. The solution to this problem includes the traditional requirements that ratios of marginal utility equal relative price ratios:

$$\frac{U_2}{U_1} = p_h(r^h)$$

and

$$\frac{U_3}{U_1} = w(r^w) \ .$$

In addition, maximization requires that

$$U_3 = \lambda w(r^w), \tag{4}$$

where λ is the Lagrangian multiplier applied to the budget constraint in the maximization problem.

This condition, and the derivative of the Lagrangian with respect to r^h, imply that

$$-p_h{}'h = (w + p_c)\, c_1 \,. \tag{5}$$

This is the classical condition for equilibrium in residential location (Muth, for example). At the optimal r^h, increases (decreases) in r^h generate increases (decreases) in commuting costs and savings (increases) in housing costs which are of equal magnitude and opposite sign.

The condition in equation 4 and the derivative of the Lagrangian with respect to r^w imply that

$$w'e = (w + p_c)\, c_2 \,.$$

This condition is the analogue to equation 5 for workplace location. At the optimal r_w, increases (decreases) in r^w generate reductions (increases) in commuting costs and in wages which are of equal magnitude and completely offsetting.

This result demonstrates the theme on which this study is based; in some sense moves and quits are 'substitutes'. Here, residence and workplace locations move in opposite directions in response to the same shock. For example, if commuting costs increase through an exogenous increase in p_c, equality in equation 5 requires reductions in r^h until the resulting reductions in c and increases in p^h balance. In this circumstance, equality in equation 6 requires increases in r^w until the resulting reductions in c and w balance.

This model illustrates the trade-offs which are essential to spatial equilibrium for worker/consumers. The marginal adjustments described by this model demonstrate the tendency for residence and workplace changes to differ in direction. However, these adjustments neglect the discontinuities which characterize actual spatial changes. Adjustments to r^h require moves. Adjustments to r^w require quits. Both actions are discrete, and subject to fixed costs.

These costs are fixed temporally; they only happen once per move or quit, regardless of the expected duration of tenure at the next residence or workplace. More importantly, in the context of this paper, they are fixed spatially. Quitting or moving entail some costs regardless of whether the new workplace or residence is next door to or many miles from the old. These costs constrain the worker/consumer's ability to alter workplace and residence locations in response to spatial disequilibria.

The budget constraint in equation 3, incorporating fixed costs, becomes

$$w(r^w)L_0 = [w(r^w) + p_c]c(r^h, r^w) + w(r^w)L + p_h(r^h)h + x + MC_M + QC_Q \,, \quad (6)$$

where C_M and C_Q represent the fixed costs of moving and quitting, respectively. $M = 1$ if $dr^h \neq 0$ (if a move occurs), 0 otherwise. $Q = 1$ if $dr^w \neq 0$ (if a quit occurs), 0 otherwise.

With this condition, the marginal analysis above is insufficient to solve the the worker/consumer's maximization problem. In effect, the worker/consumer now has four different budget constraints. Each corresponds to one of the four different choice pairs for (M, Q); (0,0), (1,0), (0,1) and (1,1). Optimization requires comparing the optimal choices under each budget constraint to determine the (M, Q) pair that yields the global utility maximum.

Intuitively, the effect of these fixed costs must be to reinforce 'substitution' of quits and moves. If a commute is suboptimal, complete adjustment through workplace or residence relocation alone incurs only one fixed cost. Adjustment through simultaneous workplace and residence relocations achieves the same goal, but incurs the fixed costs of both. This strategy can only dominate under unusual combinations of wage and housing price gradients. Ordinarily, moves or quits, individually, should be sufficient to reestablish equilibrium.

An example demonstrates this intuition. [2] For convenience, assume that housing consumption and hours of work are constant, and $c(r^h,r^w) = c(r^h - r^w)$. Assume that utility maximization requires a commute shorter then that from current workplace (r_1^w) and residence (r_1^h). Three alterna-

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[2] The interactions between quitting and moving would be more completely portrayed if explicit functions for $U(x,h,L)$, $c(r^h,r^w)$, $p_w^h(r^h)$ and $w(r^w)$ could yield explicit solutions for x, h, L, r^h and r^w in the model of equations 1 and 6. Then, utility levels under the four budget constraints could be compared directly. Unfortunately, this model does not appear to admit explicit solutions, at least with Cobb-Douglas utility. This problem appears intractable because the model is very nonlinear in r^w. Workplace location affects not only c, but w, and therefore the prices of commuting and leisure.

tive strategies accomplish this; moving to a new residence at $r_2^h < r_1^h$ without quitting, quitting to a new workplace $r_2^w > r_1^w$ without moving, and choosing some pair r_2^h and r_2^w simultaneously. Assume $r_1^h - r_2^h = r_2^w - r_1^w$.

Moving from r_1^h to r_2^h reduces commuting costs but increases per-unit housing prices. The net savings to moving, with workplace r_1^w, are:

$$[w(r_1^w) + p_c] \, [\, c(r_1^h - r_1^w) - c(r_2^h - r_1^w)] - h \, [\, p(r_2^h) - p(r_1^h) \,]$$

The net savings to moving, following a quit to workplace r_2^w, are:

$$[w(r_2^w) + p_c] \, [\, c(r_1^h - r_2^w) - c(r_2^h - r_2^w)] - h \, [\, p(r_2^h) - p(r_1^h) \,]$$

Moving is sensible only if the associated savings exceed C_M. Is this more likely when moves occur without or with quits? Savings to moving without quitting exceed those to moving and quitting if

$$\frac{w(r_1^w) + p_c}{w(r_2^w) + p_c} > \frac{c(r_1^h - r_2^w) - c(r_2^h - r_2^w)}{c(r_1^h - r_1^w) - c(r_2^h - r_1^w)} \tag{7}$$

The ratio $[w(r_1^w) + p_c]/[w(r_2^w) + p_c]$ exceeds one. If the time cost of commutes increases linearly or faster with distance $r^h - r^w$, the ratio $[c(r_1^h-r_2^w) - c(r_2^h-r_2^w)]/[c(r_1^h-r_1^w) - c(r_2^h-r_1^w)]$ is less than one. In these circumstances, moving without quitting generates larger savings than moving following a quit.

If, as assumed above, the time cost of commuting increases more slowly than distance, inequality 7 holds where the wage gradient is sufficiently 'steeper' than the commute function. This restriction is plausible, because workplaces are more centralized than residences in contemporary urban areas. If the wages in central workplaces did not more than compensate for the costs of commuting to the central city, all workers would prefer maximum values for r^h -- to take advantage of the gradient in p_h -- and $r^w = r^h$, implying, presumably, c=0. All workers would try to live and work at the city boundary.

With these considerations, moving without quitting is more likely to generate larger savings than moving following a quit. Therefore, the benefits of moving are more likely to exceed C_M, the fixed costs of moving, without quits than with. Moves are more likely to take place if quits do not.

The analysis of quits is analagous. Quitting from r_1^w to r_2^w reduces commuting costs but also reduces wages. The net savings to quitting without moving are:

$$\{[w(r_1^w) + p_c]\ c(r_1^h - r_1^w) - [w(r_2^w) + p_c]\ c(r_1^h - r_2^w)]\} - e\ [\ w(r_1^w) - w(r_2^w)\ ]$$

The net savings to quitting following a move are:

$$\{[w(r_1^w) + p_c]\ c(r_2^h - r_1^w) - [w(r_2^w) + p_c]\ c(r_2^h - r_2^w)]\} - e\ [\ w(r_1^w) - w(r_2^w)\ ]$$

in either case, quitting is sensible if these savings exceed C_Q.

Inequality 7 describes the conditions under which the savings to quits without moves exceed those to quits following moves, as well as those in which moves without quits yield higher savings than moves following quits. Furthermore, this inequality holds for increases, as well as reductions in commutes. Under the general conditions for which it holds, moves are more likely to take place without quits, and quits are more likely to take place without moves. [3]

II. Quits, Moves, Fixed Costs and Spatial Disequilibrium

An unusual 'experiment' provides an opportunity to compare the empirical relationships between quits and moves to the hypotheses above. In 1971, a firm employing nearly 800 employees and located in the central business district (CBD) of a large U.S. metropolitan area announced that it would relocate to a near suburb as of March, 1974. [4] This same firm has made available its annual payroll records for eight years, those between 1971 and 1978, inclusive. These records document employee move and quit behavior both in response to this relocation, and in periods when the workplace was fixed.

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[3] Andrulis draws similar conclusions from a model with uncertainty. Weinberg/Friedman/Mayo demonstrate the importance of fixed costs in move decisions by low income households.

[4] Both the central city and the metropolitan area are among the ten most populous in the United States.

Employee-years for white employees of this company comprise the sample analyzed here. [5] In the 3558 usable employee years, average employee age is nearly 35 years. Average tenure is nearly 8 years. Average weekly earnings are slightly above $300 in 1980 dollars.

The company payroll records record end-of-year addresses and employment status. Employees whose end-year addresses differ from one year to the next have 'moved'. [6] With this definition, moves are determinate for only employees with more than one year of tenure, observed during the seven calendar years 1972 through 1978. Employees who separated voluntarily have 'quit'.

Table 1.

Moves and Quits, Entire Sample

	No Quit	Quit	Totals
No Move	70.8%	11.1%	2914
Move	16.1%	2.1%	644
Totals	3089	469	3558

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[5] The discussion of the previous section may not apply to black employees because discrimination constrains their residence location choices (Kain/Quigley).

[6] Specifically, employees whose successive end-year addresses are in different 'Transportation Analysis Zones' have moved. The metropolitan council of governments divides the metropolitan area into approximately 1200 Transportation Analysis Zones. These zones are similar to census tracts in size. Area zip codes contain as few as one and as many as twenty. A matrix of travel times between all zone pairs, provided by the council of governments, is the only source for employee commute times. This definition of 'moves' includes only those residence relocations which can yield changes in measured commute times.

Table 1 presents, for white employees during the period 1972 through 1978, the distribution of employee-years over the four possible (M,Q) pairs discussed in the previous section. In spatial equilibrium, or with substantial fixed costs, quits and moves should be infrequent. One or the other of these conditions appears to characterize much of the sample; neither moves nor quits occurred in nearly 71% of the employee-year sample.

The discussion of the previous section predicts that, because of fixed costs in particular, either moves or quits would be more likely than both in spatial disequilibrium. Table 1 is consistent with this prediction, as well. Of those employee-years in which at least a quit or move occurred, in only 7.2% did both.

Comparisons between quit and move frequencies for those employees directly affected by the workplace relocation emphasize both the importance of spatial equilibrium and the importance of fixed moving and quitting costs. Employees who worked at the original workplace prior to 1974 were reacting to spatial disequilibria, attributable to the workplace relocation, in 1974 and 1975. These employees in these years were in 'transition' between equilibria with the CBD and suburban workplaces.

Among white employees with service in 1974 at the suburban workplace, 472 had previously worked at the CBD workplace. Of these, the 1973 residences of 211 had been closer to the new than to the old workplace by automobile.[7] This group was in disequilibrium after the relocation, but because its members had gained welfare through the reduction in their commutes. They might have further increased their welfare, for example, by moving into more distant suburbs, trading some of the commute reduction for increased housing consumption as dictated by equation 5. However, with fixed costs, many would have been content to enjoy their gains solely as additional leisure.

In contrast, the 261 white workers whose 1973 residences were more distant, by automobile, from the new than from the old workplace suffered from the relocation. If they moved, they might ultimately reap substantial welfare gains because the new workplace was suburban, more convenient to desirable housing than the old. If they quit, they might reasonably expect to reestablish their old equilibria through employment with a different company still located in the CBD. In either case, these employees had substantial incentives to reequilibrate.

Table 2 confirms these expectations. Employees whose commutes were reduced -- 'winners' -- by the relocation were ten percentage points more likely to neither move nor quit than were 'losers', whose commutes

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[7] Employee commute times are the times for automobile trips between transportation analysis zones of residence and of workplace. Company surveys indicated that approximately 90% of white workers commuted to the CBD workplace by car, even though this destination was better served by mass transit than any other in the metropolitan area. That proportion was plausibly much higher at the suburban workplace, where bus service was poor.

Table 2.

1974 Moves and Quits for Prior Employees Gaining and Losing From Workplace Relocation

	1973 Residence Closer to Suburban Than CBD Workplace			1973 Residence Closer to CBD Than Suburban Than CBD Workplaces		
	No Quit	Quit	Totals	No Quit	Quit	Totals
No Move	72.0%	8.5%	170	62.5%	16.1%	205
Move	18.0%	1.4%	41	20.3%	1.2%	56
Totals	190	21	211	216	45	261

were increased. Among winners that quit or moved, movers were twice as plentiful as quitters. Among losers, moves were slightly more frequent than among winners, but quits were nearly twice as frequent.

At the same time, table 2 is inconsistent with the pattern of marginal adjustments that would result from workplace relocation if fixed moving and quitting costs were insignificant. Despite the relocation, more than 60% of all losers made no spatial adjustments at all. More than 70% of all winners chose to take their gains solely in the form of reduced commutes.

All these employees chose to accept dramatic changes in their commuting costs, rather than incur the fixed costs of moving and quitting. The average automobile commute of losers at year-end 1974 was 27.9 minutes.

After any spatial adjustments they might have made in that year, this still exceeded their average 1973 commutes by 10.4 minutes, or 59%. For winners, the average commute at end-year 1974 was 15.4 minutes, 8.6 minutes or 36% less than the average for end-year 1973.

III. Econometric Model of Quits and Moves

The descriptive results of the last section support the suggestion that quits and moves are subject to substantial fixed costs which cause them to substitute for each other in the maintenance of spatial equilibrium. However, the simple correlation in the entire sample between quits and moves is -.026 and insignificant, because of the multitude of employee-years in which neither moves nor quits take place. Multivariate analyses, which control for many exogenous determinants of quit and move behavior, demonstrate this substitution explicitly.

Simultaneous probit models are the appropriate statistical representations for joint estimates of quit and move probabilities. [8] Simultaneous probit models allow estimates of the correlation between random shocks to move and quit propensities, and, if identified, direct estimates of the interactions between quits and moves. These models are

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[8] Weiss justifies single probit estimation for quit propensities, alone. Venti/Wise justify probit estimation in a complicated model of move propensities.

maximum likelihood estimates of the parameters β_1, β_2 and ρ in the following likelihood function:

$$L = \prod_{i=1}^{3558} \int_{b_1}^{a_1} \int_{b_2}^{a_2} \phi_2(\epsilon_1,\epsilon_2;\rho)\, d\epsilon_2\, d\epsilon_1 ,$$

where ϕ_2, the bivariate normal density function, is

$$\phi_2(\epsilon_1,\epsilon_2;\rho) =$$

$$[2\pi\sqrt{(1-\rho^2)}]^{-1} \exp\,[-(1-\rho^2)^{-1}\,(\epsilon_1^2 + \epsilon_2^2 - 2\rho\epsilon_1\epsilon_2)]$$

with $E(\epsilon_1) = E(\epsilon_2) = 0$, $Var(\epsilon_1^2) = Var(\epsilon_2^2) = 1$, and ρ as the correlation coefficient. The limits of integration depend on the pair (M,Q) as follows:

If $(M,Q)=(0,0)$, $a_1=-X_{1i}\beta_1$, $b_1=-\infty$, $a_2=-X_{2i}\beta_2$, $b_2=-\infty$.

If $(M,Q)=(1,0)$, $a_1=\infty$, $b_1=-X_{1i}\beta_1$, $a_2=-X_{2i}\beta_2$, $b_2=-\infty$.

If $(M,Q)=(0,1)$, $a_1=-X_{1i}\beta_1$, $b_1=-\infty$, $a_2=\infty$, $b_2=-X_{2i}\beta_2$.

If $(M,Q)=(1,1)$, $a_1=\infty$, $b_1=-X_{1i}\beta_1$, $a_2=\infty$, $b_2=-X_{2i}\beta_2$.

X_{1i} and X_{2i} are row vectors of exogenous variables which determine, respectively, move and quit propensities. β_1 and β_2 are the associated parameter column vectors. This model is the discrete analogue to the seemingly-unrelated-regression technique for continuous variables. [9]

X_{1i} and X_{2i} contain individual-, neighborhood- and year-specific variables, in addition to a constant. Both contain all individual-specific variables recorded in the company payroll tapes -- dummy variables for males and clerical workers, continuous variables for age, age squared, tenure, tenure squared, natural logarithms of current and past real earnings.

In this paper, automobile commuting time is a 'neighborhood' attribute defined by transportation analysis zones. X_{2i} contains the automobile time between the current residence zone and the current workplace, and the difference between current and past automobile times. X_{1i} contains only the past time, because the current time at year-end is endogenous to the choice of moving during the year.

The 1970 census tract of residence defines the neighborhood for the measurment of other neighborhood characteristics. The determinants of move propensities include characteristics which may index neighborhood amenities, stability and mode choice. These are percents of blacks in tract population, high school graduates among tract adults, tract population aged greater than five that had not moved between 1965 and

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[9] Successful estimations of this model require large samples. In consequence, estimates below represent the pooled sample of all employee-years. Individual employees vary in the number of times they enter this sample. This 'unbalanced design' renders estimation of individual-specific effects difficult, if not impossible. Therefore, these estimates disregard them.

1970, 1970 tract housing units vacant, and 1970 tract resident workers commuting to work by bus; 1969 tract median income and 1970 tract median owner-occupied housing value.

The determinants of quit propensities, X_{2i}, include characteristics which measure neighborhood income and stability. Among these are the percent of tract population aged greater than five that had not moved between 1965 and 1970, 1969 median tract income, and 1970 tract male and female unemployment rates. X_{2i} also contains the only year-specific variable; the metropolitan area unemployment rate. [10]

Model 1 in table 3 presents this specification, estimated for the entire sample. Parameter estimates are plausible: Moves are less likely with age, more likely with higher current earnings given past earnings, less likely with higher past earnings given current earnings, [11] and

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[10] Annual indexes for consumer prices and housing expenditures, experimentally included in X_{1i}, contributed nothing to model explanatory power.

[11] These coefficients imply that moves are a significant positive function of earnings growth. Coefficients on log current and past earnings estimate effects of earnings growth according to the following equation:

$$\beta_w \ln w + \beta_{w-1} \ln w_{-1} = (\beta_w + \beta_{w-1}) \ln w - \beta_{w-1} (\ln w - \ln w_{-1}) .$$

The coefficient on log previous earnings, with positive sign, is the implied coefficient on earnings growth. The estimated standard error of this coefficient is valid regardless of whether it is interpreted as the effect of log previous earnings or of earnings growth. Under the reformulation in terms of growth, the coefficient of log current earnings is positive, relatively small and marginally significant.

Table 3.

Bivariate Probit Estimates
of Quit and Move Behavior

Explanatory Variables	Model 1 Move	Model 1 Quit	Model 2 Move	Model 2 Quit
Log Likelihood	-2566.1		-2560.7	
Correlation Coefficient	-.252 (3.38)		-.218 (2.99)	
Quit	-	-	.0813 (1.06)	-
Quit During Relocation	-	-	-.122 (3.38)	-
Move	-	-	-	.335 (1.49)
Move During Relocation	-	-	-	-.298 (2.95)
Constant	1.04 (1.51)	4.39 (5.06)	.883 (1.18)	4.99 (5.63)
Male	-.227 (2.95)	.0588 (.374)	-.234 (3.00)	.152 (.842)
Age	-.0609 (3.30)	-.0659 (2.90)	-.0570 (2.99)	-.0488 (1.72)
Age Squared	.000503 (2.13)	.000571 (2.03)	.000470 (1.96)	.000416 (1.30)
Tenure	-.0101 (.779)	-.108 (5.24)	-.00568 (.378)	-.107 (5.26)
Tenure Squared	.0000877 (.212)	.00118 (1.27)	.0000163 (.0376)	.00135 (1.45)
Clerical	-.0614 (.776)	-.0757 (.677)	-.0652 (.817)	-.0695 (.617)
Log Earnings	1.88 (4.76)	-10.6 (21.2)	2.49 (3.24)	-11.3 (15.6)
Log Previous Earnings	-1.77 (4.68)	10.1 (20.6)	-2.37 (3.27)	10.7 (15.9)

less likely in neighborhoods with lower turnover.[12] Quits are less

[12] Moves are also, plausibly, unaffected by individual employment characteristics apart from earnings.

Table 3 (continued).

Current Commute Time	-	.00849 (2.16)	-	-
Change, Commute Time	-	.0112 (2.04)	-	-
Previous Commute Time	-.00769 (2.70)	-	-.00855 (3.00)	.00556 (1.45)
% Black	.00120 (.456)	-	.00129 (.490)	-
% High School Graduates	.00413 (1.23)	-	.00406 (1.20)	-
% in Same House, 1965	-.00694 (2.86)	-.00333 (1.44)	-.00663 (2.70)	-.0177 (.591)
Median Income, $1000's	-.0588 (2.31)	.0204 (1.47)	-.0596 (2.35)	.0208 (1.35)
Vacancy Rate	.0189 (2.06)	-	.0194 (2.11)	-
Median Value of Housing, $1000's	.0112 (1.70)	-	.0112 (1.72)	-
% Workers Commuting by Bus	-.0275 (5.06)	-	-.0283 (5.23)	-
Tract Male Unemployment Rate	-	.0199 (1.45)	-	.0232 (1.68)
Tract Female Unemployment Rate	-	.00508 (.170)	-	.00584 (.185)
SMSA Unemployment Rate	-	-.147 (8.80)	-	-.191 (8.81)

Parentheses contain asymptotic t-statistics.

likely with age,[13] tenure,[14] higher current earnings, lower prior

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[13] Weiss estimates a negative relationship between age and quits for new hires.

[14] Holmlund/Lang predict negative association between tenure and quits, holding compensation constant, when quitting entails fixed costs.

earnings, and metropolitan unemployment rates. [15]

The correlation coefficient in this model verifies the hypothesis that moves and quits substitute in the maintenance of spatial equilibrium. It is large, -.25, and significant at better than 1%. It implies that random shocks which encourage moves are likely to be accompanied by shocks which discourage quits, and vice versa.

The probability of a simultaneous move and quit is very sensitive to the magnitude of the correlation coefficient. The derivative of this probability with respect to the correlation coefficient is .509 . [16] This corresponds to an elasticity of the probability of a simultaneous move and quit with respect to the correlation coefficient equal to -6.5

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[15] Sex has no significant effect on quits. This result is consistent with those in earlier papers. Haber/Lamas/Green also find no gender differences in separation rates after controlling for income. Blau/Kahn draw a similar conclusion from their simulations. Meitzen asserts that female quit propensities increase with tenure while male quit propensities decrease. However, his equations also demonstrate that female quit propensities decrease markedly with age, to which male propensities are insensitive. These comparisons suggest that, as the maximum tenure in his sample is 2.5 years, his analysis may not adequately distinguish between the effects of age and tenure.

[16] Zax (1980) gives the formula for this derivative. Here, it is calculated at the estimated value for ρ, values of .741 and .966 for ϵ_1 and ϵ_2, respectively. These are reasonable values at which to calculate this derivative because, first, the bivariate normal distribution function gives the probability of (M,Q) = (1,1) to be 2.1% with these values, equal to the sample frequency. In addition, these values yield total quit and total move probabilities which are proportional to the sample frequencies. Values which reproduced the actual sample frequencies for quits and moves would be preferable, if they existed. However, no two values of ϵ_1 and ϵ_2 can simultaneously satisfy the three conditions of reproducing the sample frequencies for total moves, total quits and simultaneous moves and quits.

at the estimated correlation and the sample frequency for $(M,Q) = (1,1)$.

A simulation further demonstrates the magnitude of substitution implied by this correlation coefficient. Model 1 as estimated predicts that the average probability of both a quit and a move in observations of this sample is equal to .0177 .[17] If the spatial relationship between quits and moves was unimportant, model 1 would estimate the correlation coefficient to equal zero. Model 1 with a zero correlation predicts the average probability of simultaneous moves and quits to be .0276, 56% higher than that predicted with the estimated correlation.

This correlation coefficient demonstrates the 'weak form' hypothesis of section I; the imperatives of spatial equilibrium with fixed moving and quitting costs imply that shocks which encourage quits should discourage moves, and vice versa. The discussion of section I also implies a 'strong form' hypothesis; quits made for the purposes of establishing spatial equilibrium should discourage moves directly, and vice versa.

Model 1 does not estimate these direct effects. However, several exogenous variables appear in only the specification for X_{1i} or that for X_{2i}. Formally, these 'exclusion restrictions' permit identification of the direct effects of quits on moves, and moves on quits.

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[17] Probit estimation is not constrained to 'go through the means'. In other words, predicted probabilities at average values for the exogenous variables are not, as a rule, equal to the sample frequencies.

Quit effects on moves can be estimated by first estimating a single probit equation for quits, where the vector of explanatory variables X_i contains all exogenous variables in either X_{1i} or X_{2i}. The product of this vector and the coefficient estimates from the single probit, $X_i \gamma$, can be entered into a move equation as an 'instrumented' value for quits, Q^*. The coefficient on Q^* in this equation will consistently estimate the true effect of quits on moves. 'Instrumented' values for moves, M^*, can similarly be entered in quit equations. [18]

As many moves and quits take place for reasons unrelated to considerations of spatial equilibrium, quits and moves should not have systematic affects on each other in the sample as a whole. The specification of model 1, augmented with instrumented quits and moves among the explanatory variables, confirms this. It yields insignificant estimates of quit and move effects on moves and quits, respectively. [19]

However, 'transition' employee-years uniformly represent conditions of spatial disequilibrium. Reestablishing equilibrium should therefore have been a more important stimulus for quits and moves during this period. Model 2 in table 3 tests this proposition. It includes the instrumented values for quits and moves during the period of workplace relocation as separate variables, in addition to these variables for the

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[18] This procedure is analogous to three-stage-least-squares with continuous dependent variables. Mallar derives it formally. Maddala provides a useful summary. They recommend the use of Q^* and M^*, rather than the value of the normal distribution function associated with Q^* and M^*, in part to insure that identification does not depend solely on functional form.

[19] The author can provide estimates of this model.

whole sample. [20]

This model emphatically supports the 'strong form' hypothesis. Instrumented quits and moves for the whole sample are insignificant in the move and quit equations. In contrast, with better than 1% significance, quits during relocation reduce the probability of moves and moves during relocation reduce the probability of quits. [21]

These effects are substantial. For an employee whose probability of moving was equal to the sample frequency of .182, a transition quit would reduce that probability by 29.1%, to .129 . For an employee whose probability of quitting was equal to the sample frequency of .132, a transition move would reduce that probability by 40.2%, to .079 . [22]

Lastly, this model confirms the relationship between random shocks to

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[20] The specification of model 2 drops the variable measuring current commute time from the equation for quits. Were it included, it would also be among the instruments for move propensities. Because moves and current commute times are simultaneously determined, this would be improper.

[21] Only one other paper considers interactions between quits and moves. Weinberg estimates significant positive effects of moves on quits and quits on moves for individuals in eleven groups defined by gender, ethnicity and residential tenure. These results may derive from his sample, in which individuals do not have common workplaces, do not report incomes or tenure and may not be employed. They may also be artifacts of his statistical technique. Unfortunately, he obtains these estimates from seemingly-unrelated linear regression models in which the original dummy variables for moves and quits serve as dependent variables in one equation and explanatory variables in the other. This procedure is severely biased.

[22] The derivatives of quit and move probabilities yield similar comparisons. They are probably less useful than these simulations because moves and quits are not continuous variables.

move and quit behavior in model 1. Explicit controls for direct effects of quits and moves on each other slightly reduce the magnitude of the correlation coefficient. Nevertheless, the correlation coefficient in model 2 remains strongly and significantly negative.

Table 4.

Correlation Coefficients and Coefficients on Endogenous Quits and Moves By Period

	Model 3	Model 4	Model 5	Model 6
Quit in Move	-	.502	-.232	.423
Equation		(1.16)	(1.81)	(2.33)
Move in Quit	-	1.07	-1.25	.397
Equation		(2.46)	(2.05)	(1.04)
Correlation Coefficient:				
Period 1	-.142	-.0666	-	-
	(1.07)	(.458)		
Period 2	-.367	-	-.447	-
	(2.25)		(1.97)	
Period 3	-.269	-	-	-.234
	(2.42)			(2.06)
Log Likelihood	-2565.5	-688.1	-597.8	-1196.4
Observations	3558	949	896	1713

Parentheses contain asymptotic t-statistics.

Table 4 presents parameter estimates which reiterate the above distinction between move and quit interactions in equilibrium and disequilibrium periods. Model 3 duplicates the specification of Model 1, with the exception that it allows for correlation coefficients to differ for employer-years prior to the workplace relocation (1972 and 1973, period

1), the transition years (period 2) and the years of the stable suburban workplace (1976 through 1978, period 3).[23]

These correlation coefficients are all negative, consistent again with the hypothesis of substitution. They also provide additional evidence that quit and move interactions are particularly strong in periods of spatial dislocation. The transition correlation, -.37, is quite large. Finally, the strong negative correlation in period 3, -.27, provides evidence that shocks to quits and moves remain negatively correlated when workplaces are stable.

As measured by a likelihood ratio test, model 3 is not significantly better than model 1. Nevertheless, models 4 through 6 of table 4 validate these observations. These three models duplicate the specification of model 2 for periods 1 through 3, individually. Collectively, their explanatory power is significantly greater than that of model 2.[24] The correlation coefficients in each are similar to the three of model 3, though that of model 4, for period 1, is somewhat smaller and that of model 5, for the transition period, is even larger.

Moreover, these models confirm that the strong form substitution hypothesis holds uniquely for periods of spatial dislocation. In the

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[23] Period 3 also includes employee-years in 1975 for employees who were first hired at the suburban workplace, in 1974.

[24] The likelihood ratio test of this hypothesis yields a chi-square value of 166.4, with 68 degrees of freedom. The critical value for this test at .1% significance, with 70 degrees of freedom, is only 112.3 .

transition period, quits and moves have significant negative effects on each other that substantially exceed, in magnitude, the effects of model 2. In contrast, all effects of one on the other in periods 1 and 3 are positive.

IV. Conclusion

Moves and quits may be spatially disequilibrating for individuals with tenuous attachments to a specific job or community. For example, workers quitting to accept new jobs in other cities or regions will naturally move, as well. For them, either a quit or a move may cause disequilibrium for which the other is the solution. Such individuals appear to be rare in the sample examined here.

The intuition developed here is that, for individuals who intend to stay in the same metropolitan area, moves and quits are equilibrating. Spatial equilibrium can fail for many reasons. For example, if utility is not separable in leisure, price changes for any other consumption good will ordinarily render the current commute suboptimal. When spatial equilibrium fails, either a move or a quit is sufficient to restore it.

The empirical results strongly support the hypothesis that quits and moves substitute for each other in the maintenance of spatial equilibrium. At all times, shocks which encourage one tend to be associated with shocks which discourage the other. At times of spatial dislocation, quits and moves substitute directly; the occurence of one substantially reduces the probability of the other.

These results will not surprise multi-plant employers, who often have a policy of providing relocation bonuses for employees they transfer between plants. In the language of this paper, the transfer creates spatial disequilibrium. Relocation bonuses reduce the fixed costs of moving without changing those of quitting. Without bonuses, employees might use quits to reequilibrate. With them, moves become more likely, instead. Similarly, they will be familiar to employers who have relocated (presumably with selective or no relocation bonuses) for the purpose of enocouraging voluntary separations among unwanted employees.

The policy implications of these results derive from the recognition that quits occur for reasons of spatial equilibrium, as well as for reasons related to workplace conditions. 'Voluntary restrictions' on imports of inexpensive automobiles, the construction of a highway or a subway, the abandonment of a bus route or the institution of a substantial gasoline tax will create spatial disequilibria. The results here do not estimate the costs of these disequilibria. However, they demonstrate conclusively that worker/consumers will move or quit to reequilibrate. Particularly in 'tight' housing markets, policies which alter the 'prices' of commutes may provoke unexpected changes in job mobility.

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