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INTEREST RATE VARIATIONS, MORTGAGE PREPAYMENTS AND HOUSEHOLD MOBILITY

John M. Quigley*

Abstract—The volatility of interest rates and the deregulation of the mortgage lending sector have meant that many homeowners also own mortgages at terms more favorable than current interest rates. This paper presents a model of residential mobility decisions and an empirical analysis which evaluates the importance of the ownership of these mortgages upon the mobility of homeowners.

The results, based upon proportional and non-proportional hazard models, indicate that these effects are quite large. The empirical analysis distinguishes between different regulatory regimes which govern the assumption of existing mortgages, and indicates the implications of these findings for the pricing and valuation of mortgage backed securities.

I. Introduction

THE increased volatility of interest rates in the 1980s, coupled with the deregulation of the mortgage lending sector, had two well-documented effects upon the housing market and upon financial institutions. First, alternative methods of housing finance flourished, as buyers and sellers devised "creative" methods to finance housing transactions, often including the assumption of an existing mortgage together with a seller-financed second mortgage. Second, the imbalance in the term structure of the assets and liabilities of savings institutions meant that, with higher interest rates, the net worth of these institutions declined precipitously. Existing long-term mortgage assets declined in value, while at the same time, savings institutions were forced to pay market interest rates on liabilities, mostly short-term savings accounts. These impacts on the housing market have been carefully studied (see Jaffee (1984) for a review of the former, and Carron (1982) for a review of the latter).

These events probably had a third important impact upon consumer behavior and upon the health of savings institutions, arising from the "lock-in" effect of the ownership of mortgages at favorable terms—namely a decline in the residential mobility of homeowners and an increase in the average duration of outstanding mortgages. This "lock-in" effect is potentially important from several viewpoints. First, since almost two thirds of American households are homeowners, a small change in the average residential mobility of households, arising from ownership of below-market-rate loans, could translate into a large decrease in the inter-urban and intra-regional mobility of labor.[1] Second, any variation in the probability that mortgages will be prepaid, arising from this lock-in effect, has a direct impact on the market value of the assets of savings and loan institutions and the true net worth of these institutions (regardless of the net worth calculated by industry rules of thumb and accounting practices).

Finally, the sensitivity of the duration of mortgages to interest rate variation is of practical importance in secondary financial markets. For example, the price of collateralized mortgage obligations (CMOs) should be expected to vary with the prepayment probabilities of mortgages. If these probabilities are not constant, but have a predictable component, it follows that the pricing of CMOs will reflect systematic differences in mobility rates or prepayment rates.

This paper provides an empirical analysis of the lock-in effect of favorable mortgage terms upon the housing market. Section II below presents a simple model of the mobility behavior of households and relates the lock-in analysis to the existing literature on residential mobility and mortgage prepayment. In section III, the parameters of the model are estimated from observations on individual households, their housing consumption, and

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* University of California, Berkeley.

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[1] For example, the average mobility rate for American households, estimated from the Current Population Survey (CPS), was 20% per year for the 1970s. For 1980 through 1983, it had declined to 16.6%. Mobility, according to the CPS definition and the definition used in this paper, is defined as a change in the residence of a preexisting household.

their home mortgage terms, during a period of rising interest rates, 1979–1981. Section IV notes some of the implications of the model.

II. A Model of Moving Behavior

Households typically adjust housing consumption to desired or equilibrium levels by moving, or for homeowners, by investment in home improvements. If there were no transactions costs, a household would adjust its housing immediately whenever desired consumption changed. The costs of searching and moving, however, are substantial, and instantaneous adjustments to variations in housing demand are implausible. According to theory, in response to a change in housing demand, a household will decide to search for an alternative dwelling if the expected utility gains from the search outweigh the utility costs of searching, concluding the transaction, and moving. (See Venti and Wise, 1984, for an explicit model of housing demand and moving costs.) In the most sophisticated empirical analysis of residential mobility to date, the costs of searching and moving are enumerated empirically for a sample of low income renter households (Weinberg, Friedman and Mayo, 1981). These costs are found to exert a profound impact on residential mobility.

For homeowners, with perfect capitalization of financial terms and with assumable mortgages, there would be no significant transactions costs associated with mortgage ownership. However, empirical analysis strongly suggests that differing financial terms are not fully capitalized into housing values (see Durning and Quigley, 1985, Benjamin and Sirmans, 1987), and many mortgages are not legally assumable (discussed below). Under these circumstances, if interest rates rise, the homeowner who sells his house is not fully compensated for the value of his existing loan. This will increase the transactions costs of moving, perhaps substantially.

In this analysis, we assume that housing demand is governed by permanent income (proxied by the annual income of households, the age of the head of household, and the race and education of the head of the household) and family size; changes in the demand for housing arise from variations in permanent income, marital status and household size.[2] For homeowners with a given length of tenure, we relate the probability of moving to these factors—annual income, age, race, education, household size, increases and decreases in household size and a dummy variable representing change in the household head. We also estimate the cash value of each household's mortgage, if it holds one, and investigate the relationship between the magnitude of the lock-in and residential mobility.[3]

In a given year, the probability of moving by any homeowner is rather low, and empirical evidence strongly suggests that this probability declines with years of tenure (Quigley and Weinberg, 1977). Logically, the problem of estimating the mobility relationship is similar to that of constructing a life table for individuals of given socioeconomic characteristics. To observe directly the distribution of years of tenure at the point at which households move, one would have to wait until all individuals in a particular sample moved to tabulate the results. In the absence of the appropriate panel of individuals, the information conveyed by those who have not moved in any cross section can hardly be ignored. Indeed, the treatment of truncated observations is crucial.

The analysis reported below is based upon the so-called Cox (1972) model and its extensions. Consider a special case, the proportional hazards model (Kalbfleisch and Prentice, 1980). Define T as the length of tenure at which a household makes a move. Let f be the probability density function on T, and F be its cumulative distribution function. Let X be the vector of contemporaneous and historical characteristics, $X(t) = \{x(\tau), 0 < \tau$

[2] There is a large literature on the correlates of residential mobility, much of it undertaken by planners and demographers (see Boehm (1981) for an example drawn from a part of the same data source utilized below). The variables noted above are broadly consistent with that literature.

[3] There has been relatively little published work on mortgage prepayment, and only one study (Green and Shoven, 1986) is based upon micro data. Apparently all of these studies have been confined to the analysis of information about mortgages themselves, despite Green and Shoven's caveat "It is most important to recognize that the primary determinants of the decision to sell a house are not related to interest rate fluctuations. They are largely concerned with the personal circumstances of the owner. . . ." (p. 43). Those few aggregate studies of prepayments have regressed the prepayment proportions of mortgage pools (Arak and Goodman, 1985) or cohorts (Peters, Pinkus, and Askin, 1984) on average interest rate spreads and other time varying parameters. For an early study, see Curley and Guttentag (1974).

$< t\}$. The probability that a household with characteristics X does not move until t years of tenure is

$$1 - F(t; X) = \text{prob}(T \geq t; X). \tag{1}$$

The conditional density of moving at t, given that the household has not previously moved (the "hazard rate", λ), is

$$\lambda(t; X) = \frac{f(t; X)}{1 - F(t; X)}. \tag{2}$$

In the simplest case, assume that the hazard rate is separable and proportional and that the hazard depends only on the current value, $x(t)$ (Kalbfleisch and Prentice, 1980, Sec. 5.5),

$$\lambda(t; x) = \lambda(t)\pi(x_1, x_2), \tag{3}$$

where $\lambda(t)$ is the "base line" hazard, a function of years of tenure. The factor π, which varies with the socioeconomic characteristics of the household, x_1, and the value of the mortgage terms, x_2, at t is assumed to be of the form

$$\pi(x_1, x_2) = \exp(\beta_1 x_1 + \beta_2 x_2) \tag{4}$$

where the β's are parameters.

No parametric model is assumed for the underlying survival function, but the specification of equation (3) does imply that the ratios of the hazard functions for individuals with differing characteristics (x_1, x_2) are independent of time. The separability assumption facilitates estimation of the parameters, β and $\lambda(t)$, by permitting the log-likelihood function to be factored into two components and maximized iteratively.

More generally, it may be more reasonable to presume that the importance of mortgage terms in affecting mobility decisions is not independent of tenure time. If, as argued by Dynarski (1985), households' attachment to neighborhoods increases over time, then the same increase in lock-in or transactions costs may affect the mobility pattern of households with varying years of tenure very differently (see Ioannides, 1984). A test of the proportionality assumption is straightforward, if computationally cumbersome.

$$\begin{aligned}\lambda(t; x) &= \lambda(t)\pi(x_1, x_2, t)\\ &= \lambda(t)\exp(\beta_1 x_1 + \beta_2 x_2 + \beta_3 x_2 \log t)\\ &= \lambda(t)t^{\beta_3 x_2}\exp(\beta_1 x_1 + \beta_2 x_2). \end{aligned} \tag{5}$$

The proportionality assumption is tested by joint estimation of the parameters $\lambda(t)$ and β by maximum likelihood methods.

III. Empirical Analysis

The empirical analysis is based upon three samples of homeowners extracted from the Panel Study of Income Dynamics (PSID) which has been gathered by the Survey Research Center of the University of Michigan each year since 1968. For three years, and for only three years, questions about homeowners' mortgage terms were included in the annual survey. For 1979, 1980 and 1981 the survey asked homeowners to indicate the outstanding balance (B), if any, on the home mortgage, the number of years remaining on the mortgage (n), and the annual debt service, i.e., payment of principal and interest (A). The samples for analysis consist of all homeowners surveyed in 1979, 1980, and 1981 whose responses to these questions passed three consistency checks.[4] The samples consist of 1,768 homeowners in 1979, 1,092 in 1980, and 1,142 in 1981 (out of a total of 2,605 homeowner households sampled in 1979, 2,557 in 1980 and 2,483 in 1981).[5]

For each observation, we compute the present value of the mortgage *premium*, x_2, defined as the

[4] First, households were excluded from the analysis if the "accuracy code" for the values recorded for the remaining mortgage principal and for the annual mortgage payment signified assignment of responses by PSID interviewers.

Second, households were excluded from the analysis if the mortgage interest rate (r) implied by the recorded values of B, n, and A lay outside the range of 2% to 27%. The mortgage interest rate is defined as the value of r that satisfies:

$$B - \sum_{t=1}^{n} \frac{A}{(1+r)^t} = 0. \tag{N-1}$$

Third, households who did not move were eliminated from the sample if the interest rate computed from equation (N-1) above varied by more than 5% between adjoining years.

[5] As noted, some 33% to 58% of homeowners failed one or more of these consistency checks and are excluded from subsequent analysis. Those households reporting consistent information on mortgage terms were similar in education and income to other homeowners, but were somewhat older, on average. The age difference probably results from the increased likelihood that older homeowners have no outstanding mortgage balance:

Average Value	1979	1980	1981
Income (thousands)			
All owners	$24.1	$23.6	$23.7
Analysis sample	24.5	23.1	25.8
Education (years)			
All owners	12.2	12.3	12.4
Analysis sample	11.5	11.2	11.4
Age (years)			
All owners	46.0	45.3	44.8
Analysis sample	50.1	54.5	54.0

TABLE 1.—HAZARD RATE MODELS OF HOUSEHOLD MOBILITY[a]

	Proportional			Nonproportional		
Variable	1979	1980	1981	1979	1980	1981
Present Value of Mortgage Premium (×10,000)	−0.572 (5.06)	−1.125 (4.78)	−0.582 (3.96)	−1.370 (2.27)	−0.910 (3.42)	−1.280 (2.36)
Increase in Family Size (number)	0.402 (2.53)	0.558 (2.46)	0.841 (3.47)	0.342 (2.26)	0.577 (2.60)	0.831 (3.44)
Decrease in Family Size (number)	0.046 (0.45)	0.064 (0.25)	0.034 (2.65)	0.108 (1.90)	0.061 (0.23)	0.339 (2.63)
Income of Family (×100,000)	0.172 (0.36)	0.140 (5.15)	−0.069 (0.13)	0.027 (0.53)	0.134 (5.56)	−0.786 (0.14)
Age of Family Head (×100)	−8.988 (11.12)	−7.508 (7.06)	−7.521 (6.84)	−0.093 (12.39)	−7.293 (6.93)	−7.453 (6.79)
Education of Family Head (years)	0.155 (3.92)	0.096 (1.72)	0.073 (1.43)	0.116 (3.26)	0.097 (1.78)	0.068 (1.32)
Size of Family (number)	−0.101 (1.51)	0.040 (0.41)	0.075 (0.81)	−0.010 (0.18)	(0.022) (0.22)	0.076 (0.81)
Black Household (1 = yes)	−0.319 (1.10)	0.328 (0.38)	−0.904 (2.01)	−0.319 (1.31)	0.231 (0.22)	−0.903 (2.00)
Other Non-White Household (1 = yes)	0.744 (2.06)	0.274 (0.27)	−0.526 (0.52)	0.955 (2.90)	3.806 (0.37)	−0.530 (0.52)
Change in Family Head (1 = yes)	b	1.374 (2.04)	1.227 (2.91)	b	1.451 (2.13)	1.212 (2.87)
Z^c (×10,000)				−0.200 (1.70)	−0.550 (2.31)	−0.180 (1.26)
Likelihood Ratio (Chi-square)	316.32	212.57	182.23	394.00	260.82	228.29
Observations	1768	1092	1142	1768	1092	1142
Number of Moves	134	58	71	134	58	71

[a] Asymptotic t-ratios are in parentheses
[b] No variation in sample
[c] Z is defined as the present value of the mortgage premium times the logarithm of years of tenure

difference between the outstanding balance of the mortgage (the present value of the mortgage at the contract interest rate) and the present value of the mortgage at prevailing market interest rates r_m

$$x_2 = B - \sum_{t=1}^{n} \frac{A}{(1 + r_m)^t}. \quad (6)$$

In 1979, the mortgage premium held by the average household was worth about \$570. In 1980, this premium had increased to approximately \$800. By 1981, the average premium enjoyed by households with existing mortgages was about \$1,800 when compared to a newly written mortgage for the same term at market rates. It is worth noting that the variation in the premium is quite substantial, ranging from a low of zero for households who owned their homes free and clear[6] to a high of \$57,000 in 1979 and \$66,000 in 1981. In comparison, the average self-reported market value of houses in these samples ranged from \$42,000 in 1979 to \$45,000 in 1981.

Table 1 reports the parameters β of the proportional hazard rate models estimated by maximum likelihood for each of the three samples. For each year, the table presents the coefficients and their asymptotic t-ratios.[7] Among the demographic variables related to housing demand, increases in household size are consistently related to residential mobility. The hazard rate also declines with the age of the household head and increases with the education of the head. These findings are consistent with the great bulk of empirical evidence on household mobility. There is also weak evidence that homeowner mobility rates vary by race, but the effects of the other variables are generally inconsistent or insignificant.

In each of these "Cox-regressions," the coefficient on the variable measuring the lock-in effect of higher interest rates is highly significant, with a t-ratio of four or five. Note that these statistical models control explicitly for the age and length of tenure of households. Since most mortgages are written at the time households move

[6] Actually the minimum values are −\$120, −\$132 and −\$140, respectively, presumably reflecting the transactions costs of simply refinancing.

[7] Coefficients are normalized so that the vector sum $\beta x = 1$ where x is the vector of means of the independent variables.

into dwellings, the outstanding balance (and hence the potential lock-in effect) declines with years of tenure. Note also that each model is estimated from cross-sectional data, so the lock-in effect for each observation is measured with respect to a single prevailing market interest rate rather than a time-varying parameter. Within any year, dispersion in interest rates used to compute the mortgage premium arises only from their term structure.

Table 1 also reports the results of estimating the more general non-proportional hazard rate model. In these equations, the variable Z denotes the product of the present value of the mortgage premium and the logarithm of years of tenure—that is, Z is $\log[t^{x_2}]$ in equation (5), whose coefficient is β_3. The coefficients of most of the demographic variables are quite similar. The mortgage premium variable remains highly significant; the interaction term is less clearly significant, at least statistically. The negative sign on the coefficient implies that the importance of the lock-in effect, relative to other factors affecting decisions to move or to stay at a given location, declines as the length of tenure increases. This is consistent with the increased residential attachment from "community" enjoyed by long-term residents, at least according to one recent study (Dynarski, 1985).[8]

It should be stressed that this empirical analysis is based upon observations on the moving behavior of individual households. It may be reasonable to make inferences about household prepayment behavior, but we do not directly observe that transaction.[9] It is of course possible for households to prepay existing mortgages without moving. During periods of declining interest rates, we may expect to observe such behavior, as households refinance mortgage debt. Of more importance to a mortgage prepayment interpretation of these results, however, is the possibility that households may move without prepayment, by selling the right to assume their mortgages. We can provide some additional evidence on this point, comparing the behavior of those homeowners residing in states where "due-on-sale" clauses were enforceable with those households residing in states where these clauses were unenforceable during the period 1979–1981.[10] In states where the due-on-sale clauses typically written into mortgage contracts were enforceable, the decision to move by a homeowning household almost invariably resulted in mortgage prepayment.[11] In other states, households could move without prepaying mortgages by selling their mortgages along with their dwellings, presumably capitalizing some of the benefits.

Table 2 disaggregates the variable measuring the lock-in for homeowners living under these two kinds of institutional arrangements. As measured by the likelihood ratio statistic, the importance of the lock-in does vary under the two regulatory regimes. The absolute value of the parameter measuring the lock-in is about 50% higher for the households living in states enforcing due-on-sale clauses than those living in other states. This is to be expected, since in the former group of states there was no way for moving households to retain or capitalize any of the benefits from ownership of mortgages at more favorable terms than the current market. The t-ratio of the variable measuring lock-in is much larger for states where mortgage premiums cannot be capitalized, and these differences are statistically significant, though barely so.

The results of the more general non-proportional hazard rate models, also in table 2, are consistent with these findings, as are those of similar models, not shown, estimated separately for due-on-sale states and for states where such clauses were unenforceable.

Finally, it is possible to exploit the panel nature of the data directly to explore the importance of the lock-in and other economic variables in affecting mobility, relative to the unobserved individual-specific heterogeneity of sampled indi-

[8] As discussed below, it is also consistent with the importance of unmeasured population heterogeneity in distinguishing movers from stayers. See Flinn and Heckman (1982).

[9] This is in contrast to the recent analysis by Green and Shoven (1986) who built a multiple decrement table for California mortgages, but who had no information on those characteristics of owners that presumably underlie residential mobility.

[10] "Due-on-sale" agreements stipulate that the balance of the mortgage loan is due and payable upon sale of property. Thus an enforceable due-on-sale clause implies that homeowners may not sell the right to assume an existing mortgage. For a variety of reasons, such clauses were not enforceable in 24 states during the 1979–1981 period.

[11] The principal exception is for homeowners who moved and rented out their former residences.

TABLE 2.—HAZARD RATE MODELS OF HOUSEHOLD MOBILITY FOR STATES WITH DIFFERENT REGULATORY REGIMES[a]

Variable	Proportional			Nonproportional		
	1979	1980	1981	1979	1980	1981
Present Value of Mortgage Premium: States in Which Due-On-Sale Clauses Are Enforceable (×10,000)	−0.731 (4.13)	−1.201 (4.90)	−0.614 (3.88)	−1.640 (2.34)	−8.480 (2.65)	−4.850 (1.97)
Present Value of Mortgage Premium: States in Which Due-On-Sale Clauses Are Not Enforceable (×10,000)	−0.492 (3.47)	−0.835 (1.77)	−0.465 (1.24)	−1.060 (1.52)	−6.710 (1.43)	−1.130 (1.95)
Increase in Family Size (number)	0.391 (2.44)	0.555 (2.55)	0.837 (3.45)	0.349 (2.32)	0.561 (2.74)	0.818 (3.41)
Decrease in Family Size (number)	0.056 (0.54)	0.040 (0.16)	0.034 (2.66)	0.099 (1.71)	0.017 (0.06)	0.335 (2.57)
Income of Family (×100,000)	0.160 (0.35)	0.140 (5.21)	−0.008 (0.15)	−2.231 (0.42)	13.399 (5.53)	−0.654 (0.12)
Age of Family Head (×100)	−8.922 (11.03)	−7.588 (7.10)	−7.520 (6.81)	−0.094 (12.38)	−7.425 (1.96)	−7.207 (1.36)
Education of Family Head (years)	0.157 (3.97)	0.094 (1.67)	0.072 (1.41)	0.114 (3.22)	0.107 (1.96)	0.070 (1.36)
Size of Family (number)	−0.096 (1.43)	0.041 (0.42)	0.072 (0.83)	−0.006 (0.11)	−0.035 (0.35)	0.084 (0.89)
Black Household (1 = yes)	−0.326 (1.13)	0.310 (0.29)	−0.909 (2.02)	−0.307 (2.89)	3.861 (0.37)	−0.546 (0.54)
Other Non-White Household (1 = yes)	0.749 (2.08)	0.296 (0.29)	−0.542 (0.53)	0.957 (2.89)	3.861 (0.37)	−0.546 (0.54)
Change in Family Head (1 = yes)	b	1.393 (2.070)	1.330 2.893	b	1.544 (2.24)	1.229 (2.90)
Z^c Enforceable				−0.130 (0.84)	−7.100 (2.60)	−0.990 (1.68)
Z^c Not Enforceable				−0.250 (1.50)	2.370 (0.45)	−0.140 (0.90)
Likelihood Ratio (Chi-square)	316.95	213.93	184.48	398.15	285.34	231.34
Observations	1730	1092	1142	1730	1092	1142
Number of Moves	134	58	71	134	58	71

[a]Asymptotic t-ratios are in parentheses

[b]No variation in sample

[c]Z is defined as the present value of the mortgage premium times the logarithm of years of tenure. Z is measured separately for states in which due on sale clauses are enforceable and for those in which such clauses are unenforceable

viduals. One way to represent individual-specific heterogeneity[12] is to postulate

$$\lambda(t; x_t, \delta_t) = \lambda(t)\exp(\beta x_t + \delta_t) \quad (7)$$

where δ_t is an individual component of variation in hazard rate for individual i, and x_t is the defined covariate vector at t. Under these circumstances the parameters β can be recovered from maximum likelihood estimation of

$$\log[\lambda(t; x_t, \delta_i)/\lambda(t-1; x_{t-1}, \delta_i)] = \beta(x_t - x_{t-1}). \quad (8)$$

[12] This is surely not the only way to introduce unobserved population heterogeneity into this problem. Importantly, however, this specification makes it possible to distinguish unobserved mover–stayer propensities from observed characteristics and duration. See Flinn and Heckman (1982) and Heckman and Singer (1984).

Table 3 reports those estimates, based upon all households for which there are multiple observations in the three individual samples. For each of the multiple observations on households, the hazard function is related to the deviation of each variable from its mean value for that household. Thus all individual-specific time invariant characteristics are held constant (including measured characteristics such as race and education as well as unmeasured characteristics).

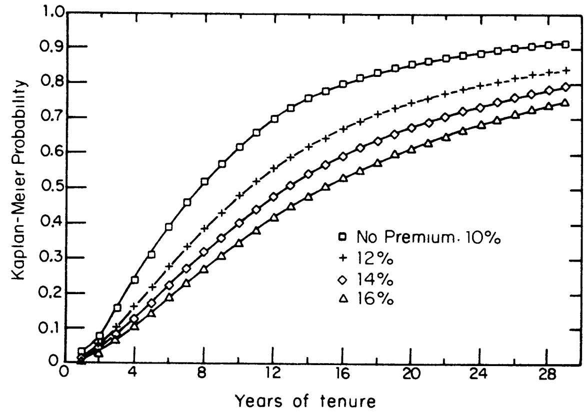
Even in this formulation, based on a very short time series and a differencing procedure which increases the noise-to-signal ratio, the hazard model as a whole is highly significant. The variable measuring the deviation of income from its mean has a t-ratio of five and the coefficient of the change in family size measure is highly significant. Most important, the coefficient of the deviation of the lock-in is highly significant.

TABLE 3.—PARAMETERS OF HAZARD RATE MOBILITY MODELS USING INDIVIDUAL DIFFERENCES[a]

Variable[b]	Proportional		Nonproportional	
Present Value of Mortgage Premium ($\times$10,000):				
All States	−0.820		−1.330	
	(2.47)		(3.44)	
Due on Sale Enforced		−1.142		−1.630
		(2.82)		(4.84)
Not Enforced		−0.710		−0.090
		(1.98)		(0.17)
Increase in Family Size (number)	0.280	2.572	0.590	0.625
	(1.25)	(1.36)	(2.85)	(2.73)
Decrease in Family Size (number)	0.654	0.711	0.585	0.588
	(3.88)	(3.92)	(5.23)	(5.22)
Income of Family ($\times$100,000)	7.772	6.952	1.764	1.796
	(5.25)	(5.16)	(2.89)	(2.99)
Change in Family Head (1 = yes)	−0.449	0.462	−0.398	0.405
	(0.79)	(0.83)	(0.84)	(0.85)
Z^c ($\times$10,000)				
All States			−0.360	
			(3.75)	
Due on Sale Enforced				−0.520
				(5.20)
Not Enforced				−0.020
				(0.11)
Likelihood Ratio (Chi-square)	69.45	70.31	52.90	67.84
Observations	2597		2597	
Number of Moves	207		207	

[a]Asymptotic t-ratios are in parentheses.
[b]All variables are measured as deviations from the mean value recorded for *each* individual.
[c]Z is defined as the present value of the mortgage premium times the logarithm of years of tenure.

FIGURE 1.—CUMULATIVE PROBABILITY OF MOVING



IV. Some Implications

The magnitude of these transactions costs in inhibiting residential mobility can be investigated by computing the baseline hazard and the variations from the hazard attributable to the value of below market mortgages. The baseline hazard $\lambda(t)$, $t = 1, 2, \ldots, 30$ is estimated by maximum likelihood, and the baseline survival probabilities are computed by the Kaplan-Meier method (see Kalbfleisch and Prentice, 1980, Sec. 4.3).

For example, from table 1, column 3, the cumulative probability of moving, P_t, for a household who moves into its dwelling at age 35, at the mean of the other variables is

$$P_t = 1 - \prod_{i=1}^{t-1} S_i^*, \tag{9}$$

where

$$S_i^* = S_i \exp[-0.0752(53.99 - \{35 + i\})]. \tag{10}$$

The term in square brackets is the proportional adjustment for a household whose head's age deviates from the sample average of 54. S_i is the Kaplan-Meier survivor probability with years of tenure i, estimated from the hazard rate, $\lambda(i; x)$.

Figure 1 presents a comparison of the mobility pattern of a homeowner who, at age 35, takes out a $40,000 fixed rate, 10% mortgage with a thirty-year repayment period at several different market interest rates. If interest rates remain at 10%, the probability of moving climbs to one-half after about eight years of tenure. This is not far from the industry rule of thumb (Struck, 1974). If interest rates climb to 12%, the lock-in for this household is $5,800 after the first year, $5,000 after the seventh year, and almost $4,100 after the twelfth year. At a 16% rate, the lock-in is $13,700 after the first year, and almost $10,300 after the twelfth year. The decrease in mobility over the life cycle, attributable to these variations in market conditions, shown in figure 1, is quite striking.

In a competitive market, the differences in consumer behavior which underlie the simulations in figure 1 will be reflected in the prices and values of mortgage-backed securities. Table 4 reports the relationship among the portfolio values implicit in the consumer mobility patterns reflected in figure 1. The first row presents the value of a pool of fixed 30-year 10% annuities as a fraction of face value, computed at various interest rates. As the table indicates, with no prepayments a 10% mortgage drops in present value and market price by one fourth as interest rates move to 14%. Most mortgages permit prepayment with little penalty, and presumably this call option is priced along

TABLE 4.—VALUE OF A MORTGAGE POOL: TEN PERCENT, 30-YEAR, FIXED RATE, $40,000 MORTGAGES AT VARIOUS INTEREST RATES, 1981

	Interest Rate			
	10%	12%	14%	16%
No Prepayment	1.000	0.853	0.741	0.653
Average Prepayment Experience[a]	1.000	0.867	0.764	0.683
Including Lock-in Effect[b]	1.000	0.862	0.755	0.668
Including Lock-in Effect and[c] Non Proportional Hazard	1.000	0.859	0.749	0.660

[a] From estimates of baseline hazard, λ, for purchaser at age 35, at the mean values of other variables.
[b] Including estimated effect of favorable mortgage terms on mobility, from table 1, column 3.
[c] Effect of mortgage terms on mobility estimated from table 1, column 6.

with other aspects of the mortgage. Given the determinants of residential mobility and the average prepayment experience of the households in this sample, but ignoring the lock-in effect, the second line of the table computes the present values of all cash flows. These entries are computed for a household moving into a dwelling at age 35 at the mean of the other characteristics. Rows three and four indicate how the value of mortgage assets decline as the lock-in reduces residential mobility and hence prepayments. The additional decline in portfolios attributable only to increased duration varies from 1% to 3.4% as interest rates increase from 12% to 16%.

Taken together, these results emphasize the importance of the mortgage market in affecting the residential mobility of households. They also suggest that better models of consumer moving and payment behavior can greatly increase our understanding of mortgage markets and financial institutions.

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