The Transition to Homeownership:
The Importance of Early Career Concerns

Stephen Cameron
Department of Economics
Columbia University
New York, NY 10027

Joseph Tracy
Domestic Research
Federal Reserve Bank of NY
New York, NY 10045

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PRELIMINARY

The aggregate homeownership rate in the United States declined slightly between 1980 and 1990, following three decades of increases. This raised the issue of whether housing affordability had become a problem in the 1980s. Green (1996) examines several measures of affordability applied to housing markets defined at the state level. He finds no significant evidence of general affordability problems.

The aggregate homeownership numbers mask a more fundamental change in housing during the last decade. Figure 1a shows age/ housing tenure profiles for all heads of households based on the 1970, 1980, and 1990 Census of Population data. Comparing the 1980 age/tenure profile to the 1970 profile we see that homeownership rates for young households less than thirty-two years old were virtually unchanged over the decade. For households older than thirty-two, the 1980 profile reflects uniformly a higher ownership rate than a decade earlier. In sharp contrast, over the decade of the 1980s the age/tenure profile significantly steepened. Homeownership rates fell for younger households (less than fifty years in age) and increased for older households. For households with heads aged thirty the ownership rate declined by eight percentage points. The Chicago Title & Trust survey data on first-time home buyers also indicates an increase over the decade in the average age of first home purchase from 28.1 to 30 years.

While this steepening of the age/tenure profile has received some attention in the literature [Gyourko and Linneman (1993), Linneman and Megbolugbe (1992), and Wachter (1990)] , we feel that it remains an open question. Three candidate explanations are demographic changes, affordability problems, and increased household mobility. In this paper, we focus on labor market developments for young workers in the 1980s and the role they might have played in the decline in homeownership for young households. Specifically, we explore how changes in real wages and job stability early in workers' careers have affected the transition to homeownership.

We limit our attention to the first transition to ownership. This is the starting point for a more complete analysis. We use the National Longitudinal Survey of Youth (NLSY) data to estimate the determinants of the rent-to-buy transition. We argue that it is important to use panel data rather than cross-section data to measure these transition probabilities. We simplify the analysis by conditioning on school completion and household formation [see Haurin, Hendershott and Kim (1994)].

What Do We Know About the Steepening of the Age/Tenure Profile?

There are at least three possible explanations for why homeownership rates declined for young households during the 1980s: demographic changes, affordability problems, and increased mobility. The prime demographic factors for the transition to ownership are the age of first marriage and household composition. Affordability issues relate both to the user cost of housing and to the income prospects of households. We will examine mobility primarily from the standpoint of the changing stability of jobs for young workers.

The age of first marriage continued its upward trend during 1970s through the 1980s. From 1980 to 1988 the average age of first marriage for men increased from 23.6 to 25.5, while for women the average age increased from 21.8 to 23.7.[1] Another trend is the increase in the share of young female-headed households. From 1980 to 1990 the percent female-headed households increased from 12 to 14 percent for whites, 40 to 44 percent for blacks, and 20 to 24 percents for hispanics.[2] While both of these factors would be expected to delay the transition into ownership, it is unlikely that they provide the sole explanation for the steepening of the age/tenure profile. Apagar *et al* (1990) table A-12 report homeownership rates from 1973 to 1989 disaggregated by marital status and the presence of children. For households aged 25-34, the data indicate declines in ownership for all groups except *single with no children*. In addition, figure 1b shows the age/tenure profile restricted to male-headed households. This shows the same steepening as figure 1a which is based on all households.

Many researchers have raised affordability as the most likely cause of the steepening age/tenure profile. Wachter (1990) comments on the decline in real incomes for low-income, young, and less educated households over the 1980s. Gyourko and Linneman (1993) argue that housing prices for low to moderate quality houses (those likely purchased by first-time buyers) did not decline in real terms over the 1980s, leading to likely affordability problems for younger and less-educated households. Megbolugbe and Linneman (1993) summarize their interpretation of the data as follows.

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[1] *Statistical Abstract of the U.S.* (1995). table 145.

[2] *Statistical Abstract of the U.S. (1996).* table 74.

> "But not only are home ownership rates among the young low, they also fell throughout the 1980s. Binding wealth constraints for the young rather than increased mobility probably explain this phenomenon. Together, transaction costs and wealth constraints severely discourage home ownership among the young." [page 671]

While affordability issues are likely an important part of the explanation, we feel that the potential role for increased mobility has not been adequately explored in the literature. The affordability hypothesis suggests that the declines in homeownership rates among young households should be concentrated among the lower income and less well educated. Table 1 disaggregates the changes in homeownership rates for young households by income quartiles and by years of schooling. The table shows declines across all income quartiles and levels of educational attainment.

Further evidence pointing away from affordability comes from the Chicago Title & Trust Company data on first-time home buyers. While the average age of first-time home buyers increased by almost two years over the 1980s, the average years that households save for a downpayment decreased during the middle of the decade. Similarly, the average percentage of the downpayment coming from gifts declined slightly over this period. Mayer and Engelhardt (1996) argue that gifts appear to be a method of relaxing a binding wealth constraint. Taken together, this evidence suggests that young households may be deferring the time that they begin to prepare to actively save for their first house. One reason for deferring the purchase decision may be increased mobility.

We view mobility for young households as primarily arising out of early career concerns. Topel and Ward (1992) find significant turnover of jobs among young workers for several years until a more permanent job match is found. In the first ten years in the labor market, they find that a typical worker will hold seven jobs (about two-thirds of his/her lifetime total). The job turnover probability declines rapidly with a workers accumulation of tenure on the job.

This intensive period of job shopping and turnover for young workers plays an important role in maximizing the returns to their human capital investments. The first ten years of a career will account for roughly two-thirds of the lifetime wage growth for a male high school graduate [Murphy and Welch (1992)]. Wage gains at job changes accounts for at least a third of early-

career wage growth [Topel and Ward (1992)]. Not surprisingly, the wage is a key determinant of a young worker's mobility. Controlling for individual heterogeneity, Topel and Ward find that a 10 percent within-career increase in the wage reduces the probability of changing jobs by about 20 percent, and jobs with more rapid wage growth are also more stable.

Job shopping by young workers is also a key determinant of their residential mobility. Bartel (1979) calculates the fraction of geographic moves that are "caused" by a worker's decision to change jobs.[3] She estimates that the fraction of geographic moves initiated because of job changes was 47% for young workers and 33% for older workers.[4] Similarly, the *Fannie Mae National Housing Survey* in 1994 asked a sample of owners and renters their estimate of the probability of moving in the next two years. The average probability was 14 percent for owners and 33 percent for renters.

To update Bartel's findings we turned to data on recent movers from the 1989 *American Housing Survey*. Question 52 asks recent movers what are the reasons they moved from their last residence. They are also asked to indicate the main reason. We classify a move as "caused" by career concerns if they indicated either *New Job or Job Transfer* or *Other, Financial/Employment Related* as the primary reason indicated. The remaining responses can be broadly categorized as quality of life reasons, demographic reasons, and affordability.[5] Unfortunately, one of the response categories is *Change From Owner to Renter OR Renter to Owner*. For movers who indicated this as the primary reason, we reallocated their response equally among the other reasons indicated. If no other reason was indicated, then we deleted them from the sample.

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[3]Geographic moves involve either a change in county or a change in SMSA. Thus, all intra-urban moves are excluded from her calculations.

[4]The estimates for young workers is derived from the National Longitudinal Survey of Young Men and covers workers age 19-29 in 1971 and moves occurring between 1971 and 1973. The estimates for older workers is derived from the National Longitudinal Survey of Mature Men and covers workers age 45-59 in 1966 and moves occurring between 1966 and 1971.

[5]Examples of quality-of-life reasons include, *Wanted Better Quality House* and *Other Housing Related*. Demographic reasons include *Needed Larger House or Apartment, Married, Widowed, Divorced or Separated,* and *To Establish Own Household*. Affordability reasons include *Wanted Lower Rent or Less Expensive House to Maintain*.

Table 2 presents the distribution of main reasons given for the recent move by age, education, and distance of the move. Overall, roughly one-fifth of all residential changes are caused by job-related reasons. This fraction is constant across our two broad age groups. Distance as measured by whether the move is intra-urban or inter-urban is a determinant of the importance of job changes for residential mobility. For young workers only 7% of intra-urban moves are job driven, whereas roughly 23% of within state inter-urban moves and 64% of between-state inter-urban moves are job driven. There are no clear differences by age groups in these findings. However, persistent differences do emerge by education level. For each age group and each distance of move, the fraction of residential changes caused by job-related reasons increase with the level of education of the worker. These findings conform closely with Bartel's earlier findings.

The importance of job mobility in overall mobility for the NLSY data is illustrated in table 3. The table gives cross-tabs on 26,507 person year observations drawn from the data. For each person year observation we ask whether a job change occurred during the year and whether a change of state residence occurred during the year. The high rate of job mobility is indicated by the fact that over 61% of the observations involve a job change. Residential mobility is much lower with 16% of the observations involving a change in MSA residence. Job changes are clearly important in residential changes with nearly 82% of residential changes also involving a job change.

There is growing evidence that the process of workers sorting into eventual long-term jobs may have changed during the 1980s. Farber (1995) and Sweinnerton and Wial (1995) find evidence of a decrease in job stability for young and less educated workers. This shows up both in the distribution of current tenure and in job retention probabilities. Diebold, Neumark, and Polsky (1996) raise several estimation issues concerning the Sweinnerton and Wail study, leaving the specific evidence on job stability somewhat still in question.

A Housing-Transition Model

This section lays out several methodological issues concerning the estimation of the transition into homeownership. As we are interested in explaining the shape of the aggregate

age-homeownership profile of individuals early in the life-cycle, we specify a model that can account for age-by-age determinants of the transition into homeownership. Three issues are discussed in this section. First, some notation is developed and a simple discrete-time duration model of the transition to homeownership is laid out. We distinguish the age-by-age homeownership transition probabilities (the *flow* into ownership) from the age-by-age homeownership state probabilities (the *stock* of owners). Models based on state probabilities are difficult to interpret and do not isolate the right parameters to understand the *transition* from renting to buying. Second, we discuss why past research that has focused on estimating state probabilities — sometimes because better data were not available — may have misstated the relative effects of demographic and labor market forces as well as homeownership costs on the decision to buy. Third, we compare our approach to the continuous-time duration models used by a number of other researchers of the rent-to-own transition.

We are interested in modeling the transition from renting to homeownership. To begin, we assume a simple model in which people may be in one of two housing states: renting or owning. Let r_t^i represent whether household i rents at age t and let o_t^i denote ownership at age t. To keep the notation at a minimum, we suppress the superscript i unless necessary. At each age an individual who rents at age $t-1$, moves into ownership at age t if the value of all present and future benefits of homeownership relative to renting outweigh the costs; that is if

$$V_{o_t, r_{t-1}} - V_{r_t, r_{t-1}} \geq 0, \tag{1}$$

where $V_{\cdot, r_{t-1}}$ is the expected indirect utility of either owning (o_t) or renting (r_t) next year given that the individual is currently a renter. Homeowners carry out a similar calculation each period to determine whether they will continue owning or transit back to renting.

Let s_t denote the state at age t (either o_t or r_t). We assume that $V_{s_t, s_{t-1}}$ is known to be or can be approximated by a linear function.

$$V_{s_t, s_{t-1}} = \beta_{s_t, s_{t-1}} X_{s_t, s_{t-1}} + \epsilon_{s_t, s_{t-1}} , \quad (2)$$

where $\beta_{s_t, s_{t-1}}$ is a parameter vector, $X_{s_t, s_{t-1}}$ is a vector of covariates, and $\epsilon_{s_t, s_{t-1}}$ is an error term. It is important to note that the parameter vector may depend on both age (to capture age effects) and the current state (to capture state-dependence). The error term and the covariates are also allowed to depend on both age and state. We do not impose any restrictions on the error term or the parameters that arise from the intertemporal optimization problem, so this model should be considered as a reduced form.

We now specify the statistical model. To keep notation to a minimum, we let X_t represent an augmented covariate vector that includes all the variables in each $X_{s_t, s_{t-1}}$ for all possible states. Let $P(o_t \mid X_t , \ldots , X_0)$ denote the *state* probability of homeownership at age *t*. It is important to note that the state probabilities depend on current and past realizations of X_t. If an individual is not an owner at age *t*, then he rents with probability $P(r_t \mid X_t , \ldots , X_0) = 1 - P(o_t \mid X_t , \ldots , X_0)$. Together, let the vector of state probabilities for households who face covariates X_t be given by

$$P_t = \begin{bmatrix} P(o_t|X_t, \ldots X_0) \\ P(r_t|X_t, \ldots X_0) \end{bmatrix} . \quad (3)$$

Starting from a given age *t*, an individual may make the transition to any other state. Let the full transition probability matrix be represented by

$$T_{t,t-1} = \begin{bmatrix} P(o_t|o_{t-1}, X_t) & P(o_t|r_{t-1}, X_t) \\ P(r_t|o_{t-1}, X_t) & P(r_t|r_{t-1}, X_t) \end{bmatrix} . \quad (4)$$

Note that we assume here that conditional on the prior ownership state the current transition probabilities depend only on the current realization of X_t. After the transition probabilities are specified for all $t = 1, \ldots, T$, where T is the maximum age, the entire life-cycle of state probabilities are determined by the model once an initial condition is specified for age 0.

$$P_0 = \begin{bmatrix} P(o_0|X_0) \\ P(r_0|X_0) \end{bmatrix} . \tag{5}$$

Starting from an age 0 and an initial set of state probabilities, the current state probability vector for a person who faces the profile of exogenous variables $\{X_t\}_{t=0}^{T}$ is given by

$$\begin{aligned} P_t &= T_{t,t-1} \cdot P_{t-1} \\ &= T_{t,t-1} \cdots T_{1,0} \cdot P_0 \ . \end{aligned} \tag{6}$$

Thus, the probability of owning at age t is the sum of the new owners and old owners (who bought before age t) less the old owners who transit back to renting.

As specified, the current model is a first-order Markov transition model. However, the variables X_t can be easily augmented to accommodate higher-order state dependence or various types of duration dependence [see Heckman (1981) for a menu of possibilities].

In the analysis in this paper, we focus only on the *first* transition from renting to owning. To understand how this transition fits into our framework, assume for the moment that owners never move back to renting; that is, $P(r_t \mid o_{t-1} ; X_t) = 0$ for all t. Then the probability of moving from renting to owning at age t is given by

$$\begin{aligned} &P(o_t|r_{t-1},\cdots,r_0;X_t,\cdots,X_0) \\ &= \left\{1 - P(r_t|r_{t-1},X_t)\right\} \cdot \left\{\prod_{j=1}^{t-1} P(r_j|r_{j-1},X_j)\right\} \cdot P(r_0|X_0) \ . \end{aligned} \tag{7}$$

That is, with no transitions from owning back to renting the probability of buying at age t is simply the probability of being a renter at age 0 and continuing to rent until age $t - 1$, when the transition to homeownership is made. Conditional on the sequence $\{X_j\}_{j=0}^{t}$, equation (7) defines a discrete-time duration distribution (also known as a geometric distribution) of the time from age

0 to the time of first purchase. This distribution is fully determined once the transition probabilities are specified.

The probability of owning a home *by* age *t* (buy at age *t* or before) is just the sum of the probabilities defined by (7) — that is, the chances of being an owner at the initial age 0 or of buying at or before age *t*:

$$P(o_t|X_t,X_{t-1},\cdots,X_0) = 1 - P(r_o|X_0) + \sum_{j=1}^{t}\left\{\left[1 - P(r_j|r_{j-1},X_j)\right]\cdot\left[\prod_{k=1}^{j-1} P(r_k|r_{k-1},X_k)\right]\cdot P(r_0|X_0)\right\} . \tag{8}$$

In this paper we assume that each $\epsilon_{o_t,r_{t-1}}$ of equation (2) is distributed normally and that utilities are fully described by (2). Hence, the transition probability from rent to rent and the initial condition for the probability of renting take the form

$$\begin{aligned}\Phi(\beta_t X_t) &= P(r_t|r_{t-1};X_t)\\ \Phi(\beta_0 X_0) &= P(r_0|X_0) ,\end{aligned} \tag{9}$$

where (Since we have assumed thus far that owners do not move back to renting, we can replace the $\beta_{s_t,s_{t-1}}$ of equation (2) with the less cumbersome notation β_t with no loss in generality at this point.) Substituting (9) into (7), the probability of moving from renting to owning at age *t* (equation (7)) becomes

$$\begin{aligned}&P(o_t|r_t,\cdots,r_0;X_t,\cdots X_0)\\ &\quad = \left[1 - \Phi(\beta_t X_t)\right]\cdot\Phi(\beta_{t-1}X_{t-1})\cdots\Phi(\beta_0 X_0) .\end{aligned} \tag{10}$$

and the probability of owning at a given age (equation (8)) becomes

$$P(o_t|X_t,X_{t-1},\cdots,X_0) = 1 - \Phi(\beta_0 X_0) + \sum_{j=1}^{t}\left\{\left[1-\Phi(\beta_t X_t)\right]\cdot\left[\prod_{k=1}^{j-1}\Phi(\beta_{k-1}X_{k-1})\right]\cdot\Phi(\beta_0 X_0)\right\} . \quad (11)$$

Our analysis in this paper is based on the likelihood given in equation (10). Given estimates of the parameters, predicted state probabilities can then be calculated using equation (11) (see below for a more detailed discussion).

How Does Our Analysis Compare with One Based on State-Probability Estimation?

Under normality, the correct state probability is given by equation (11). Studies that estimate state probabilities from cross-sectional data or that confine their analysis to one age or one year of a panel data approximate (11). In addition, it is common to specify the probability as a logit or a probit and to assume that the covariates enter the probability in a linear fashion. Under normality, a typical likelihood would be composed of elements such as

$$\Phi(\gamma_t X_t) = P(o_t|X_t) , \quad (12)$$

where the probability being estimated is conditioned on the current set of covariates only.

Two points are worth making. First, suppose for a moment that the covariate vector of individual and housing market characteristics is unchanging over time: $X_t = X$ for all t. Then (12) could be used to estimate the true probability given by (11), though the parameters of (12) would combine estimates of the true parameters $\{\beta_j\}_{j=0}^{t}$ in an uninterpretable fashion. In addition, any age effects in $\{\beta_j\}_{j=0}^{t}$ would not be recoverable.

Second, if X_j varies by age, how should one chose the X_j so that (12) can best approximate (11)? Typically, in a cross-sectional data set an analyst has access only to X_t, though, if the information were available, one could imagine using X_0 or some function of the X_j process such as the mean of X_j from age 0 to t. The proper treatment of time- or age-varying covariates is an issue that has long been recognized in the duration literature. Though there are

few formal theorems that predict the direction of the bias, a number of studies have documented large biases when arbitrary solutions are used to approximate age- and time-varying covariate vectors [see for example Heckman and Borjas (1980)].

Both of the problems just described are compounded if, instead of a single age, an analysis is based on a sample of individuals of different ages from data collected in a given year. By focusing on age-specific transition probabilities, our study overcomes both of these limitations.

Returning to the basic model described in equations (2) through (6), how does this analysis change if we remove the assumption that $P(r_t \mid o_{t-1}, X_t)=0$ and let owners make the transition back into renting? We can generalize the simple two-state model described above by expanding the state space defined in equation (3) to distinguish between renters who are previous owners and renters who have never owned. That is, the state space now consists of four mutually exclusive categories: renters who have never owned, owners in their first spell of ownership, renters who have previously owned, and owners who have gone from renting to owning, to renting, and finally back to owning. Such a distinction is important if renters who were previous owners behave differently than renters who have never owned, due to differential tastes for housing, tax treatment of capital gains on the previous house, or some other unobserved characteristic. Under reasonable assumptions, the distinction between renter/owner and renter/owner who has previously owned is testable (occurrence dependence) as are other forms of state-dependence. In future work, we will formally determine whether such a distinction is important in our data.

For the questions addressed in this paper, we follow our earlier discussion and ignore all transitions but the first transition from renting to owning. Econometrically, no problems arise in this model by focusing only on one of many possible transitions (see below). However, because we are ignoring a sizable portion of the overall population, our decision to initially focus only on renters who have never owned may impede our ability to understand the aggregate age-homeownership profile. A quick check of the NLSY data reveals that between ages 24 and 34, five to ten percent of owners move from owning back to renting at any given age. Again, we leave the estimation of these other transitions to future work.

If these other states and their corresponding transition probabilities are important, then it is worth noting that the approximation of the true state probability $P(o_t \mid X_t, X_{t-1}, \cdots, X_0)$ by equation (12) will become even more arbitrary. Rather than the simple sum of the chances of buying by a given age as expressed in (11), the true state probability will consist of all rent-ownership *paths* that end in ownership at age *t*. That is, an individual may move back and forth between renting and owning any number of times, and equation (12) will be used to approximate the complicated underlying probabilities.

Comparisons to Previous Work

An alternative methodology to the one presented here is based upon standard duration models and builds a likelihood around the *time* an individual spends renting. This strategy is exemplified by the work of Henderson and Ioannides (1989), Gronberg and Reed (1992), and others. Boehm (1981) and Borsch-Supan (1986) are examples of analysis of housing choices based on the transition model. In general, models of stock/flow behavior can work with either the transition probability or the duration until first purchase. This can be easily seen in our case from equation (7), in which the discrete-time duration distribution is a function of all the transition probabilities.

Our preference for transition models over duration models in this application is based upon several considerations. First, the model presented here is based on transition probabilities estimated by standard logits or probits. These models are much easier to interpret than duration models as they are built around simple concepts of utility maximization. Duration models are particularly difficult to interpret when the covariates vary across time or across age. Second, age-effects — when the parameter vector β is allowed to depend on age in arbitrary ways — are more easily parameterized and understood in a transition model. Third, as we are interested in the age-homeownership profile, transition probabilities are a more natural object of interpretation than duration distributions or hazard rates.

Another advantage of this framework is that endogenous schooling, marriage, job changing and other processes can be easily accommodated by expanding the state space. We will deal with these in future work. For now, we do allow for individual-specific heterogeneity which induces correlation in the error terms of the transition probabilities across ages.

Empirical Specification and Results

We use the NLSY panel data to estimate the determinants of the first transition to homeownership. The NLSY data cover young (aged 14-21 in 1979) individuals who are making their first housing decisions during the 1980s and early 1990s. The data is relatively clean with little attrition. One drawback to the NLSY is that detailed wealth questions were not added until 1985.

We condition on individuals who have finished their schooling and who have formed an independent household. We select all males from the random sample starting at the age of twenty-four who are initially renting. Individuals who have already made the transition to ownership by twenty-four, then, are excluded from the analysis. By starting off at age 24, we attempt to circumvent problems with people who are still in school. In the NLSY, all but a small fraction of males report being out of school after age 24. We follow the individual until he makes his first house purchase, or to the end of the sample. At the time of the first purchase, the individual is dropped from the sample. As discussed earlier, in this paper we do not analyze the duration of tenancy in the first house nor subsequent housing transitions. The econometric model outlined earlier can easily accommodate these extensions. The sample, then, consists of male heads of households aged 24+ who are currently renting and who are followed until either their first purchase or the end of the sample.

The estimation proceeds in two stages. First, we use the panel nature of the NLSY data to estimate geographic mobility propensities for each individual. Second, rent-to-buy transition probabilities are estimated using the estimated geographic mobility propensities as a control variable. We use the estimated transition probabilities to simulate the state probability of owning a buying a house by age thirty. In addition, we directly estimate a state probability model and compare the coefficient estimates to simulated effects based on the transition probability model.

Geographic Mobility

Housing economists have long understood the potentially important role that expected geographic mobility plays in the decision to buy a house [see Shelton (1968) and Boehm (1981)]. The significant transactions cost in buying and selling a house discourage owning for

households who do not have a strong attachment to a community. The NLSY data do not contain any direct questions regarding a household's expected geographic mobility. Our strategy, then, is to use the panel nature of the data to back out expected mobility estimates.

For each household, we look three years in the future and ask if the household continues to live in their current state of residence. We use state rather than SMSA as our definition of geographic mobility since we are interested in the interaction between the household's choice of a labor market and a housing market. Intra-urban moves do not require that the worker change labor markets. We fit probability models to explain the observed pattern of future mobility using a set of currently observed covariates. These probability models are then used to construct expected geographic mobility propensities conditional on current values of the covariates.

We split the sample according to whether the household head is currently working full-time or not. For the non full-time employed households, we estimate a simple probit model using the indicator for future geographic mobility. The explanatory variables are two indicators for ethnic background, three schooling indicators, marital status, age, local unemployment rate, number of prior county moves in the previous three years, tenure (length of stay) in the current county, and county tenure squared.

For the full-time employed households, we estimate a seemingly unrelated probit (SUP) model on the indicators for whether the individual works for the same employer three years in the future, and for whether the individual lives in the same state three years in the future. The job mobility specification uses the control variables listed above with the number of employers in the prior three years as well as current job tenure and tenure squared in place of the prior geographic mobility measure and county tenure variables. In addition, the quality of the current job is proxied by the current log wage and hours worked. The geographic mobility specification builds on the specification for the non full-time employed sample by adding the two job quality variables as well as the predicted three-year job retention probability.

The estimation is carried out as follows. We first estimate a simple probit model for the job retention indicator using the control variables listed above. We use the estimates from this simple probit to generate predicted three-year job retention probabilities. We then estimate the SUP model using this estimate for the expected job retention probability. The estimates from this

first SUP model are used to recalculate the expected job retention probability. We use this new variable in a final estimation of the SUP model.

Table 4 provides the estimates from these job and geographic mobility models. The data generally do not indicate significant differences across ethnic groups in either job or geographic mobility. The education effects are measured relative to a high school graduate who has no college education. Individuals with some college have significantly higher job mobility rates (that is a lower probability of being with the same employer in three years). College graduates appear to have significantly higher geographic mobility. Better jobs as measured by both the current wage and hours worked are more durable in that they have a higher probability of lasting an additional three years. In contrast, controlling for the predicted job retention probability, job quality appears to have no direct effect on geographic mobility. Tighter employment conditions in the current local labor market tend to lead to both reduced job and geographic mobility. Controlling for other factors, marital status and age do not appear to be significant determinants of mobility in general. Current job and county tenure strongly affect the job and geographic retention rates. Conditional on current tenure, prior mobility as measured either by the number of jobs or county moves in the past three years also has a strong effect on future mobility. This suggests that there is heterogeneity across households in their underlying mobility patterns that is not being picked up by our other controls.

Finally, we find that the predicted three-year job retention probability has a significant positive impact on the three-year geographic retention rate. That is, households who based on their observable characteristics are more likely to continue with their present employers, are also more likely to continue residing in same state. We use the prior job mobility measure and current job tenure variables to identify this effect.

To evaluate the relative importance of job retention rates for geographic mobility we calculated for each observation in the data marginal effects associated with the job retention rate and the combined marginal effect for all other control variables. The marginal effect for the job retention rate is calculated as the change in the probability of remaining in the same state of residence for three years when we zero out the job retention variable from when we add in the job retention variable. Similarly, the combined marginal effect for all other control variables is the change in the geographic mobility rate when we zero out all other control variables from

when we add them into the specification. The average absolute marginal effect for the job retention rate is 0.02, while the average absolute marginal effect for all other control variables is 0.057. The standard deviation of absolute marginal effects for the job retention rate and all other control variables is 0.017 and 0.027 respectively. Expected job mobility, then, helps explain an important degree of the overall variation in the expected geographic mobility rates across households.

The Transition to Homeownership

In this section, we present our estimates of the first transition from renting to owning. As in our mobility specifications, we include several demographic/background variables in our specification. We control for the current marital status of the household head.[6] The number of children in the household and the change in the number of children over the past year are also included.[7] We have indicators for household heads that are black and hispanic. The education of the household head is controlled for by a series of indicators for less than high school, some college, and a college degree. Three region indicators are used to control for differential transition rates to ownership across broad regions.

We attempt to capture affordability differences across households by measuring differences in households earning capacity, as well as spatial differences in housing costs. The ability of a household to afford to purchase a house is measured by an indicator for whether the head is currently employed, his real hourly wage if employed, and his spouse's real income if he is married.

Housing costs are captured by several variables. We include a simplified version of the current user cost of ownership, $C = (1 - \tau) \cdot i \cdot P$, where τ is the household's marginal tax rate, i is the prevailing mortgage interest rate, and P is the price of housing in the local housing market. We measure P by starting with Coldwell data on comparable house prices across metropolitan

[6]We also tried specifications which controlled for the duration of the current marriage. We found no significant effect of years since marriage on the probability of a transition to ownership.

[7]We have made no attempt at this point to deal with endogeneity issues relating to marriage and children.

areas in 1986.[8] We use Freddie Mac repeat-sale house price indices to extrapolate to the remaining years in our sample. We measure the prevailing mortgage interest rate using the FHA average rate for the year. Marginal tax rates for a household are estimated by STATA, and reflect the marginal federal tax rate based on the year and household income.[9]

A limitation of the current user cost formulation is that is does not capture future changes in the user cost to the extent that they may be anticipated. In addition, only the level and not the uncertainty over the user cost is assumed to be important for the transition decision. Rosen, Rosen and Holtz-Eakin (1984) argue that uncertainty over user costs should also be relevant. They find that greater uncertainty over the relative price of owning reduces the proportion of owners.

We include two simple measures to test for whether current user costs are a sufficient statistic for housing costs. We add the mean and variability of local real house price appreciation measured over the previous three years. Higher mean appreciation rates if expected to continue should lower the user cost through the implied capital gains. In addition, given the significant transactions costs involved in purchasing and selling a house, higher expected appreciation should lower the minimal holding period required to recoup these costs.[10] Higher variability in house price appreciation is a crude proxy for greater house price risk which should discourage ownership for renters. First-time home buyers convert most of their liquid assets into down payment (and closing costs) leaving the household highly invested in housing. Declines in house prices can wipe out their equity making it difficult for the household to either refinance or move.[11] Being locked into a particular labor market can limit the households future income prospects.

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[8]This data is designed to capture relative house price differences in housing appropriate for a mid-level manager. We assume this accurately captures *relative* house price differences for lower quality housing as well.

[9]We would like to generalize this marginal rate to include single as well as joint filing status, number of dependents, and state income taxes.

[10]Higher past housing appreciation rates, though, could lead to lower transitions to ownership by making it more difficult for renters to save the required down payment, Engelhardt (1996).

[11]See Archer, Ling, and McGill (1996), Caplin, Freeman, and Tracy (1993), and Peristiani, Bennett *et al* (1996) for evidence on refinancing, and Chan (1996) on housing mobility.

There is some debate in the literature over whether user costs of housing is the correct measure of the incentive to move from renting to owning, with an alternative measure being the relative price of owning to renting.[12] The Coldwell data for 1986 includes both a house price and an equivalent rent. We use this to form a relative price of owning to renting. Unfortunately, we have not found a metropolitan rental index which would allow us to capture variation in relative rents through time. As such, our relative price measure only picks up the cross-section variation as of 1986, whereas our user cost also captures variation across time. As a result, our horse race between these two alternative measures is not entirely fair.

The last aspect we try to capture is household mobility. We use two alternative measures of expected mobility. The first is an estimate of the probability that the worker will be on the same job three years in the future. The second is the probability that the worker will be in the same state three years in the future, where this is allowed to depend on the household head's expected job mobility if currently employed full-time. These are estimated using the NLSY data for both renters and owners.

Our transition model estimates are summarized in table 5. Summary statistics on all variables are given in appendix table A1. We jointly estimate a model for the initial ownership condition at age twenty-four and the housing transitions. We estimate the model with and without allowing for unobserved heterogeneity (both between the initial condition and the housing transitions and within the housing transitions for the same household).

We will devote much of our discussion to the transition estimates rather than the initial condition estimates. Turning first to the demographic variables, hispanics and blacks have lower transition rates from renting to owning than whites with similar characteristics. Individuals who do not finish high school are significantly less likely to move into ownership all else constant. Educational attainment beyond high school appears to have no further direct effect on rent-to-buy transition rates.[13] We find a large positive and significant impact of marriage on the likelihood that a household will switch from renting to owning. Households with more children

[12] See for example Jones (1995).

[13]Higher education levels can still have important indirect effects through its impact on wages and job/geographic mobility.

as well as households who recently experienced an increase in the number of children are more likely to transit from renting to owning. These family size effects, though, are imprecisely measured. The control variables do an adequate job of explaining differences across regions in the transition rates into homeownership.

Turning to the affordability measures, we find that earnings measures have the predictable strong effect on the likelihood of a purchase. Household heads who are employed full-time are more likely to purchase a home. Similarly, higher wages by the household head and higher income of the spouse both lead to higher transition rates.

Our simple user cost measure performs quite well in capturing housing cost differences with the expected impact on the transition probability to owning. The data prefers the user cost formulation to the alternative of using only the variability in house prices. Controlling for the user cost, we find no significant effect associated with changes in the relative price of owning to renting. Consistent with Rosen, Rosen, and Holtz-Eakin (1984), we find that holding constant the current user cost transitions into ownership are higher in housing markets that have experienced faster house price appreciation rates and lower in housing markets that have experienced more variable housing prices. Both of these effects, though, are imprecisely measured.

Expected geographic mobility is an important determinant of the decision to buy a house. Our proxy for expected geographic mobility is the predicted probability that the household will remain in the same state for the next three years. Holding other factors constant, households that have a higher expectation of remaining in their current state are significantly more likely to switch from renting to owning. This finding mirrors those by Boehm (1981) based on the PSID data.[14]

In specifications (3) and (4) of table 5 we allow for a discrete distribution of unobserved heterogeneity both in the initial conditions and in the panel of transitions. While the factor loading coefficient is allowed to differ across the two models, we currently impose a common set

[14]Boehm uses two different proxies for expected geographic mobility. The first is an indicator for whether the household in fact moved over the next two years. This is in contrast to our use of the predicted probability that the household moves. The second is the household's response to the question of whether they expected to move over the next few years. Boehm reports similar qualitative findings across these two different proxies.

of probabilities. The data indicate a small percentage of the population (around 6 percent) who are significantly more likely to initially own at age twenty-four, but conditional on initially *not* owning they are significantly less likely to make the subsequent transition from renting to owning. In fact, the magnitude of the factor loading coefficient in the transition model suggests that conditional on not initially owning a house this group essentially never makes a future home purchase. Controlling for this unobserved heterogeneity has a larger impact on the coefficients in our initial conditions specification than in our transition specification.

To put these results into some perspective, we carry out some simulations that show the impact of some of our control variables on the probability that a household owns by age thirty (conditional on renting at age twenty-four). Recall that the homeownership probability can be expressed in terms of the age-specific transition rates from renting to owning. Table 6 presents these simulations.

Our three models (without random effects) all generate similar ownership probabilities of around 62% by age thirty. If we increase the user cost facing a household each year by 25%, the estimates imply a decline in the probability of owning by over 2%. Increasing the homeowner's wage by 25% leads to a increase in the probability of owning of around 2.7%. Giving a workers an additional year of tenure would increase the homeownership rate at age thirty by 2.8%. Similarly, increasing the job retention probability by 25% would lead to an increase in the ownership probability by 2.7%, an effect that is larger in magnitude than the user cost effect.

Table 6b presents comparisons of the sensitivity of several of our determinants of homeownership derived from a state probability model. That is, we took our sample of male head of households who were renting at age twenty-four and code an indicator for whether they owned a house at age thirty.[15] We estimated a probit model on this indicator using the same control variables as in our transition probability model. We see that the impact of a 25% increase in our user cost of housing measure results in no significant change in predicted homeownership rates. Similarly, the impact of a 25% wage increase is less than half of the implied impact from the transition probability model. While the impact of job mobility is similar across the two

[15] Note that this sample is not a single cross-section from the panel at a fixed date. Rather, the sample covers a range of years due to the age variation at the outset of the NLSY in 1979.

models, the state probability model produces an effect from geographic mobility that is four times the magnitude of the effect produced from the transition probability model. These results suggest that the transition probability estimates for variables such as the user cost that have considerable time-series variation for a household head are difficult to properly estimate using the state probability framework.

Conclusions

The aggregate age/housing tenure profile steepened between 1980 and 1990 with homeownership rates falling for households aged less than fifty, and homeownership rates rising for older households. In this paper, we begin investigating the possible link between job mobility related to early career development and the transition from renting to owning. We estimate a transition model from renting to the first home purchase using NSLY data on young households. We find that predicted job mobility does have a significant impact on the probability of a household being an owner by age 30. To the extent that workers entering the labor market in the 1980s can expect a higher rate of job changing early in their career as compared to earlier cohorts, our findings would imply that this feature of the labor market would delay their decision to purchase their first home.

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Figure 1a. Homeownership Profiles for All Head – of – Households.

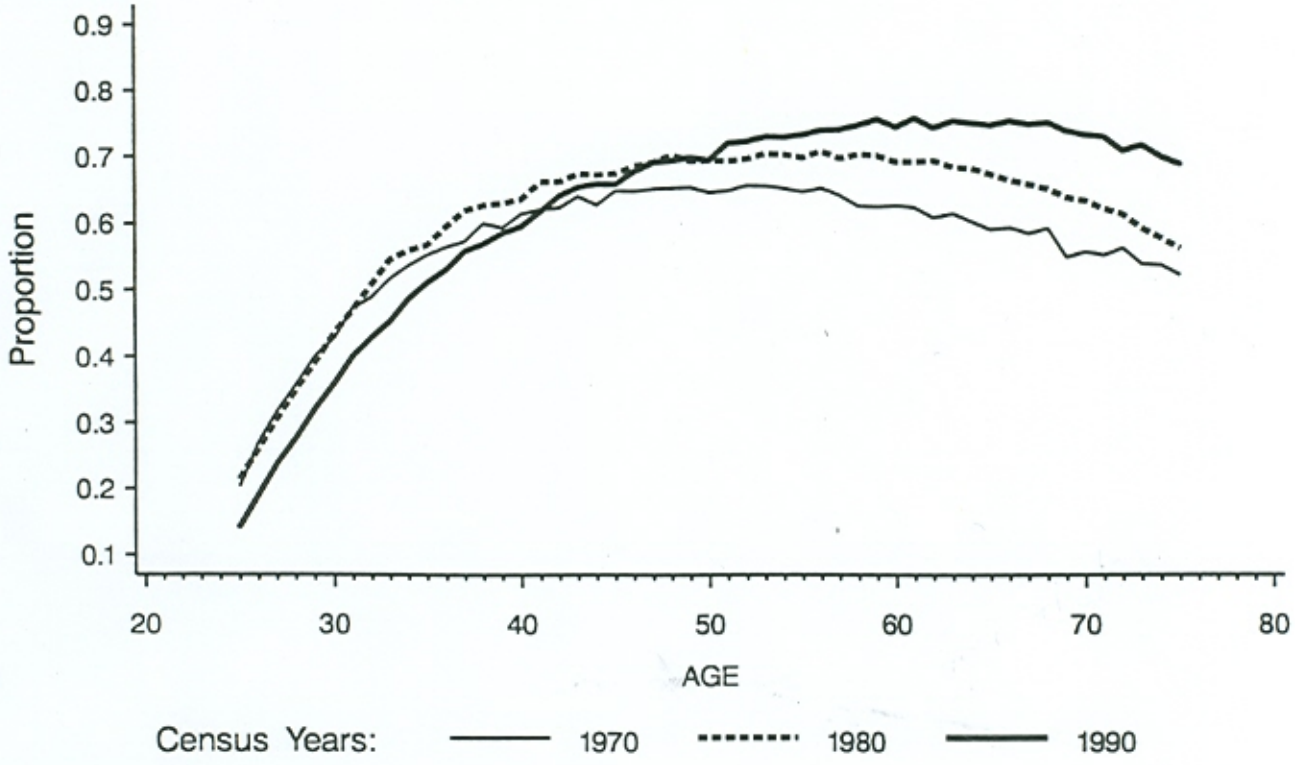
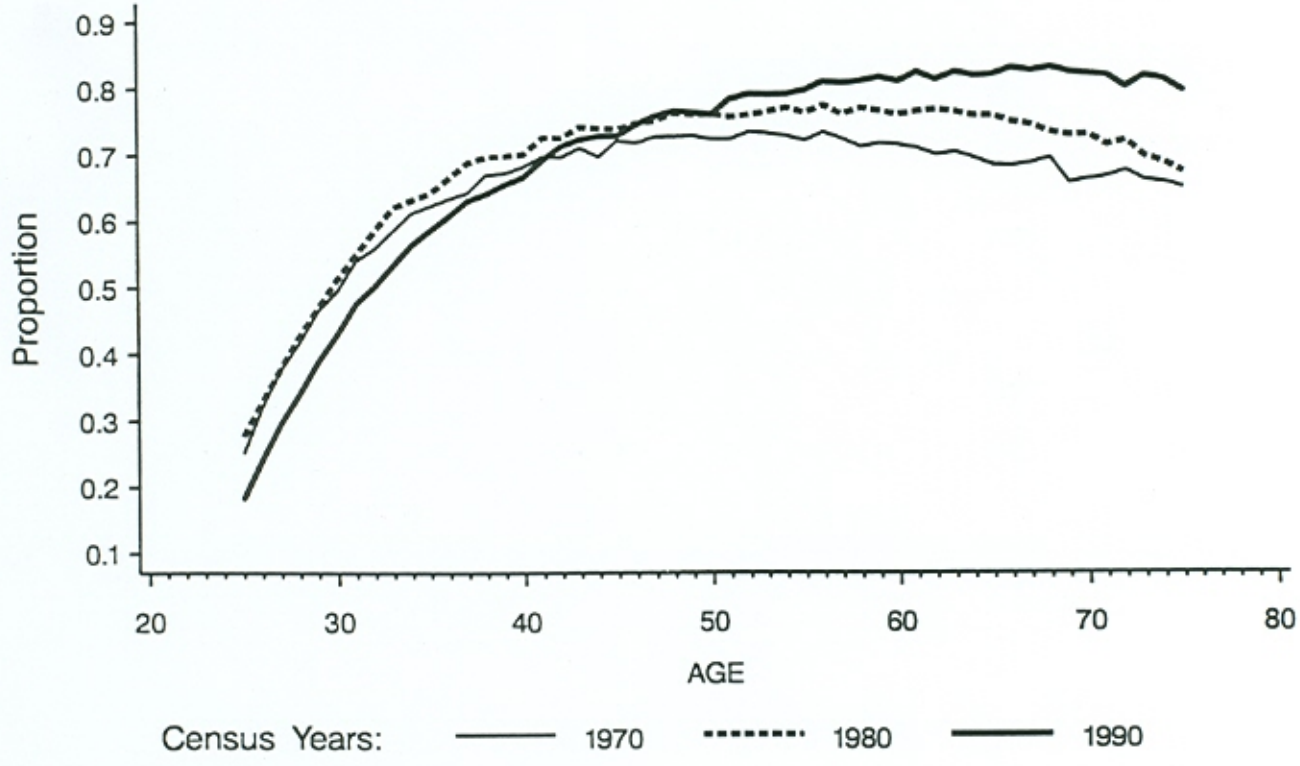


Figure 1b. Homeownership Profiles for Males.



Source: Calculations based on the 1970, 1980, and 1990 Census public use extracts.

Table 1. Change in Homeownership Rates Between 1980 and 1990

A. By income quartiles for all heads of household.				
Age	Botton 25%	25 - 50%	50 - 75%	Top 25%
25 - 29	-0.01	-0.018	-0.064	-0.087
30 - 34	-0.009	-0.048	-0.066	-0.056
35 - 39	-0.02	-0.04	-0.05	-0.021
B. By educational attainment for all heads of household.				
	Dropout	High School	13 - 15	16+
25 - 29	-0.029	-0.026	-0.052	-0.048
30 - 34	-0.057	-0.035	-0.052	-0.075
35 - 39	-0.072	-0.051	-0.05	-0.032
C. By income quartiles for all males.				
	Bottom 25%	25 - 50%	50 - 75%	Top 25%
25 - 29	-0.037	-0.06	-0.107	-0.11
30 - 34	-0.067	-0.103	-0.097	-0.056
35 - 39	-0.078	-0.067	-0.061	-0.021
D. By educational attainment for all males				
	Dropout	High School	13 - 15	16+
25 - 29	-0.058	-0.082	-0.081	-0.081
30 - 34	-0.087	-0.088	-0.065	-0.052
35 - 39	-0.114	-0.087	-0.064	-0.043

Notes: Authors' calculations using the 1980 and 1990 Census public use samples A and B. Figures represent the change in ownership rates between 1980 and 1990. "Head of household" includes both males and female heads. "All males" includes males regardless of their relationship to the household head.

Table 2. Percent of Job-Related Residential Moves

Type of Move	Education	Age Range	
		25 - 35	36 - 55
Overall	All	20.0	21.0
Intra-Urban	All	7.0	8.9
	≤ High School	6.2	9.0
	> High School	7.8	7.3
Inter-Urban, Within State	All	22.7	25.3
	≤ High School	11.9	19.0
	> High School	35.9	32.8
Inter-Urban, Between States	All	64.0	60.1
	≤ High School	49.3	47.7
	> High School	71.2	66.7

Source: 1989 *American Housing Survey*. Percents reflect the number of households that list as the primary reason for a residence change in question 52 one of the following two responses:

□ New job or job transfer

□ Other, financial/employment related

Table 3. Job and Residential Mobility - Three Year Window

Frequency Percent Row Percent Column Percent	Residential Move to Different MSA	No Residential Move	Total
Different Employer	3,527	12,712	16,239
	13.31	47.96	61.26
	21.72	78.28	
	81.70	57.29	
Same Employer	790	9,478	10,268
	2.98	35.76	38.74
	7.69	92.31	
	18.30	42.71	
Total	4,317	22,190	26,507
	16.29	83.71	100.00

Notes: Authors' calculations from the NLSY data. Sample restricted to ??

Table 4: Job and Residential Mobility

	Full-Time Employed Sample				Part-Time Employed / Non-Working Sample
Variable	Job Retention	State Retention	Job Retention	State Retention	State Retention
Constant	−1.337**	1.113**	−1.811**	0.926**	1.148**
	(0.207)	(0.311)	(0.220)	(0.331)	(0.190)
Hispanic	0.049	0.104	0.057	0.217**	0.106**
	(0.043)	(0.080)	(0.043)	(0.078)	(0.050)
Black	−0.013	0.040	−0.007	−0.023	−0.068*
	(0.042)	(0.063)	(0.042)	(0.067)	(0.041)
High School Dropout	−0.136**	0.089	−0.131**	0.077	0.063
	(0.041)	(0.069)	(0.041)	(0.073)	(0.046)
Some College	−0.122**	−0.060	−0.123**	−0.049	−0.110**
	(0.043)	(0.069)	(0.043)	(0.070)	(0.046)
College Degree	0.022	−0.215**	0.012	−0.345**	−0.365**
	(0.060)	(0.079)	(0.060)	(0.082)	(0.045)
Log(Current Wage)	0.426**	0.015	0.414**	−0.065	
	(0.039)	(0.062)	(0.040)	(0.062)	
Hours	0.005**	−0.003	0.005**	−0.002	
	(0.002)	(0.003)	(0.002)	(0.003)	
Local Unemployment Rate	0.018**	0.016**	0.014**	0.000	0.009*
	(0.005)	(0.007)	(0.005)	(0.008)	(0.005)
Married	0.043	0.091	0.038	0.056	0.066*
	(0.036)	(0.058)	(0.035)	(0.054)	(0.035)
Age	−0.011	−0.003	0.007	0.003	0.024**
	(0.008)	(0.112)	(0.009)	(0.014)	(0.007)
Number of Jobs in Past Three Years	−0.397**		-0.402**		
	(0.016)		(0.017)		
Current Job Tenure	0.263**		0.265**		
	(0.018)		(0.018)		
Job Tenure Squared	−0.014**		-0.014**		
	(0.002)		(0.002)		
Number of Counties in Past Three Years		−0.193**		-0.366**	−1.052**
		(0.035)		(0.030)	(0.021)
Current County Tenure		0.097**		0.085**	0.044**
		(0.012)		(0.012)	(0.008)
County Tenure Squared		−0.003**		−0.002**	−0.009**
		(0.0004)		(0.0004)	(0.003)
Predicted 3-Year Job Retention Probability		0.570**		0.724**	
		(0.146)		(0.146)	
Rho	0.356**		0.385**		
	(0.035)		(0.032)		
Year Effects Included	No		Yes		No

Notes: Standard errors in parentheses. ** denotes significant at the 0.05 level, * denotes significant at the 0.10 level.

Table 5: Initial Ownership Condition and Rent-to-Buy Transitions

Variable	Initial Ownership Condition	Rent-to-Buy Transition	Initial Ownership Condition	Rent-to-Buy Transition
Hispanic		−0.146 (1.49)		−0.156 (1.51)
Black	−0.251 (1.00)	−0.191 (2.04)	−0.821 (1.05)	−0.202 (2.10)
High School Dropout	0.002 (0.01)	−0.406 (3.48)	1.147 (1.14)	−0.423 (3.58)
Some College	0.451 (2.06)	−0.145 (1.47)	2.449 (2.67)	−0.137 (1.34)
College Graduate	0.274 (1.11)	0.026 (0.27)	3.137 (4.06)	0.018 (0.17)
Full-Time Working		0.493 (2.68)		0.496 (2.65)
Log(Current Wage)	0.016 (1.54)	0.017 (3.66)	−0.000 (0.01)	0.018 (3.57)
Spouse Income	0.035 (2.36)	0.022 (5.26)	0.059 (2.23)	0.024 (5.08)
Married	0.298 (1.06)	0.377 (3.31)	1.669 (2.17)	0.369 (3.12)
Number of Kids	−0.004 (0.02)	0.065 (1.33)	0.550 (1.52)	0.071 (1.39)
Change in Kids	0.063 (0.23)	0.074 (0.93)	−0.220 (0.25)	0.076 (0.93)
User-Cost of Housing	−0.001 (0.63)	−0.002 (2.22)	−0.008 (1.52)	−0.002 (2.24)
Relative Cost of Owning to Renting	−0.003 (0.56)	−0.002 (0.98)	−0.024 (1.90)	−0.002 (0.78)
Mean Past House Price Appreciation	3.688 (1.67)	0.752 (0.77)	19.933 (3.41)	0.607 (0.61)
Std. Dev. Past House Price Appreciation	3.656 (0.85)	−1.686 (0.77)	4.661 (0.41)	−1.731 (0.77)
Predicted 3-Year State Retention Probability	2.122 (1.91)	1.401 (3.55)	12.742 (7.03)	1.406 (3.48)
Intercept	−3.253 (2.50)	−2.738 (5.69)	−13.093	−2.772 (5.58)
Factor Loading			6.483	−4.890 (0.0001)
Probability Mass Point			0.938	

Notes: Specifications contain 3 region and 7 age indicators. t-statistics in parentheses.

Table 6: Model Simulations. Percentage point changes in the probability of homeownership at age 30

	Transition Model no Random Effects			Transition Model with Random Effects		
	1	2	3	1	2	3
Age 30 Predicted Probability	62.3%	62.2%	62.6%	64.0%	64.0%	64.1%
User cost up 25%	−0.021	−0.024	−0.024	−0.015	−0.017	−0.017
Own-to-rent price ratio up 25%	−0.079	−0.072	−0.071	−0.085	−0.090	−0.090
Wage up 25%	0.027	0.027	0.028	0.035	0.036	0.035
Spouse's income up 25%	0.018	0.018	0.018	0.017	0.017	0.016
Median house price change up 25%	0.003	0.003	0.003	0.003	0.003	0.003
Std Dev of price change up 25%	−0.002	−0.002	−0.002	−0.002	−0.002	−0.002
Prob. of no job change up 25%	0.027			0.032		
Prob. of no location change up 25%		0.074			0.076	
Tenure up 25%			0.021			0.030
Tenure up 1 year at each age			0.028			0.040
Indirect effect of tenure up 1 year on the ownership probability operating through:						
Prob. of no job change	0.003			0.004		
Prob. of no state change		0.017			0.025	

Notes: Simulations are based on the estimates presented in Table 5

Table 6b: Model Simulations. Percentage point changes in the probability of homeownership at age 30 from a State probability model

	State Model		
	1	2	3
Age 30 predicted probability	62.3%	62.3%	62.3%
User cost up 25%	0.003	0.004	0.004
Own-to-rent price ratio up 25%	−0.054	−0.061	−0.060
Wage up 25%	0.008	0.014	0.011
Spouse's income up 25%	0.019	0.020	0.021
Median house price change up 25%	−0.000	−0.000	−0.000
Std dev of price change up 25%	−0.018	−0.017	−0.019
Prob of no job change up 25%	0.020		
Prob of no location change up 25%		0.314	
Tenure up 25%			0.011
Tenure up 1 year			0.013
Indirect effects of tenure up 1 year on the ownership probability operating through:			
Prob of no job change	0.014		
Prob of no state change		0.017	

Note: Simulations are based on estimates from a state (stock) model of homeownership. All variables are measured at age 30

Appendix

The main source of data used are from the National Longitudinal Survey of Youth (NLSY). A cohort of young people, were we focus upon young men, are interviewed each year from 1979 to 1993. The questions are detailed and varied allowing researchers to address separate questions like homeownership and job tenure. For those familiar with the NLSY, only the random sample is used (the NLSY has a poverty and minority oversample).

A great many of the variables detailed in Appendix Table A1 are directly from the NLSY. Appendix table A1 details the final sample of 24 year olds. The table also offers some information about the variables created. An indication of which models use which information is given. Also, the table details if variables vary over time, people or location. Below are some details as to how the information was created.

Each year, respondents are asked if they own a home. Since individuals are interviewed annually from an early age to their 30s, one can determine at what age they make the transition into homeownership.

The NLSY also offers detailed information about their current employer, from which one can calculate job tenure. There are many ways to do this. Because the NLSY follows individuals, a researcher can follow through multiple interviews the job history of these young men. The basic idea behind the measure used within this paper is the length of continuous primary employment. During each interview, respondents report when they started work with their employer. Also, each survey includes a question asking if their primary employer is the same as in their last interview. If they respond yes to this, tenure is updated by linking the current job to the previous primary job. If no, the job is coded as new and the start date of the job is taken as the later date of the last interview or the time the current job started. From this, a detailed understanding of job duration can be calculated.

To control for individual heterogeneity in the probability of changing jobs, we calculate how many job changes a worker has experienced in the last three years. For those who are interviewed in a current year and the preceding three waves of the NLSY, we know if they changed employers between these yearly interviews. By adding up the number of years where each worker experiences a job change, we create a variable that captures to some extent the recent prior mobility of a worker. Analysis using the work history files of the NLSY would allow for a more comprehensive measure of past job changes and will be explored in the future.

The NLSY also specifies the location of the household. First, the region in which one lives can be obtained. For many individuals, Standard Metropolitan Statistical Area (SMSA) codes further classify the respondent's location. This is true for about 40% of the entire sample. These respondents can be linked to area specific characteristics.

Because information about current location can be obtained from the NLSY Geocode data, one can calculate how many county moves a respondent made in a fixed time period. Though not included in this version of Appendix Table A1, the number of between county moves that are recorded between the current and the last three interviews is used to control for individual heterogeneity in the probability of moving between locations. Specifically, the NLSY Geocode data allows one to know if the respondent's current county differs from the county they lived in at the time of the last interview. This is not a perfect measure as one can not discern how many times a respondent moved or if a person moved more than once and ended up in the same county that they were in the previous interview. In order to calculate past location mobility, only

individuals who are continuously interviewed remain in the sample. We also count the number of continuous years a respondent has lived in their current county. The inclusion of these measures reduces the number of observations by about one-third to one-half of the number shown in Appendix Table A1.

Two sources of housing data were used to generate a variety of housing related factors important to the decision to buy a home. First, data from Coldwell Banker were used to assign an SMSA level mean house price and mean cost of renting to a variety of SMSAs. Data for owning costs and renting costs from Coldwell banker are only available in 1986. They estimate the mean house price and rental cost for a unit of fixed quality between various regions.

To get time variation in housing costs, SMSA level house price data were obtained from Freddie Mac through their local area constant quality repeat sales price indexes. The mean change in local house prices and the standard error of the mean change in local house prices can be created directly from the Feddie Mac data.

Linking the Freddie Mac data to the Coldwell Banker data by SMSA gives us mean ownership and renter costs for about 60 SMSAs from the late 1970s until 1994. These ownership and rental costs are then merged with the NLSY observations by year and SMSA. One can then assign an ownership to renter cost variable to those members of the NLSY which live in an area covered by both Freddie Mac and Coldwell Banker data.

We also create a standard measure of user cost. The final method used was created by first obtaining family income in the past calendar year from the NLSY. From this, one can calculate the marginal tax rate. Currently, the data reflect only marginal tax rates as if all the males' family income were taxed as if they were a married family. Given the marginal tax rate, the mean local cost of housing and a yearly 30-year fixed rate mortgage rate from the Federal Housing Administration, a user cost measure can be created as (1-marginal tax rate)*(Interest rate)*(house price).

In order to correct potential measurement problems, some of the variables were truncated. Tenure was truncated at 10 years. Wage rates were truncated at 80 1994 dollars. Spousal income was truncated at $100,000 1994 dollars. The change in the number of children was truncated at -2 and at 2. Whenever any of these boundaries were reached, the boundary level replaced the reported level. In terms of the reported results, the only major change was that the coefficients on the wage rate in the transition into ownership were not statistically significant before they were truncated (there were a few large and unlikely true outliers).

Appendix Table A1 - Data Details

Name	Mean[1]	s.d.	min.	max.	N	Model[2]			Variation[3]			Source
						House	Job	Move	Time	Regional	Individ.	
Owner next year	0.11	0.31	0	1	3536	Y	Y	Y	Yes	No	Yes	NLSY
Black	0.16	0.36	0	1	3536	Y	Y	Y	No	No	Yes	NLSY
Hispanic	0.09	0.28	0	1	3536	Y	Y	Y	No	No	Yes	NLSY
Real hourly wage ($1994)	11.92	7.12	0.14	80	3536	Y	N	N	Yes	No	Yes	NLSY
North east	0.15	0.36	0	1	3536	Y	N	N	Yes	No	Yes	NLSY
South	0.32	0.47	0	1	3536	Y	N	N	Yes	No	Yes	NLSY
West	0.32	0.47	0	1	3536	Y	N	N	Yes	No	Yes	NLSY
Currently employed	0.91	0.29	0	1	3536	Y	Y	Y	Yes	No	Yes	NLSY
No high school	0.18	0.39	0	1	3536	Y	N	N	No	No	Yes	NLSY
Some college	0.23	0.42	0	1	3536	Y	N	N	No	No	Yes	NLSY
College degree	0.31	0.46	0	1	3536	Y	N	N	No	No	Yes	NLSY
Currently married	0.31	0.46	0	1	3536	Y	N	N	Yes	No	Yes	NLSY
Spouses real income (in 1000s) ($1994)	3.81	8.40	0	58.70	3536	Y	N	N	Yes	No	Yes	NLSY
Own to rent price ratio	139.49	18.48	103.33	208.38	3536	Y	N	N	Yes	Yes	No	Coldwell & Freddie Mac
Usercost measure	158.28	77.06	54.59	517.93	3536	Y	N	N	Yes	Yes	Yes	NLSY, FHA & STATA
Mean local real house price change in past 3 yrs	0.01	0.05	-0.14	0.17	3536	Y	N	N	Yes	Yes	No	Freddie Mac
s.e. in local real house price change in past 3 yrs	0.04	0.02	0.01	0.14	3536	Y	N	N	Yes	Yes	No	Freddie Mac
Local unemployment rate	6.27	2.49	2.10	18.40	3536	N	Y	Y	Yes	Yes	No	NLSY
Number of own children present	0.35	0.79	0	5	3536	Y	N	N	Yes	No	Yes	NLSY
Change in num. of children	0.60	0.38	-2	2	3536	Y	N	N	Yes	No	Yes	NLSY
Prob. of not changing Jobs in 3 yrs	0.42	0.21	0	0.90	3536	S	N	N	Yes	No	Yes	Generated by authors
Prob. of not moving states in 3 yrs	0.88	0.07	0.67	0.99	3536	S	N	N	Yes	No	Yes	Generated by authors
Current job tenure (in yrs)	2.05	2.22	0	10	3536	S	Y	Y	Yes	No	Yes	NLSY

1. Reported statistics represent values for the final random sample used in transition models.
2. Model indicates which variables were used in which model. 'House' are the models reflecting transition into ownership. 'Job' models job stability. 'Move' models locational mobility. S - some.
3. Time, regional and individual variation details if the data varies over time, region or individual.