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Estimating Equilibrium Housing Demand for "Stayers"[1]

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This study models demand by owner-occupied housing "stayers." Most consumers do not move routinely in response to small changes in income or housing price, so the "own–rent" and "move–stay" decisions are modeled as multiperiod optimization in the presence of transaction costs. The empirical section uses the "American Housing Survey" to provide a panel of household stayers for a metropolitan area. Results indicate that income and value-rent measures in different years have separable and significant impacts on housing demand. Estimated conditional income elasticities are between 0.40 and 0.45.

Key Words: housing demand; tenure choice.

1. HOUSING DEMAND WITH TRANSACTION COSTS

The economics of housing demand have evolved substantially over the past four decades. Major improvements in measuring price and income terms, the joint modeling of tenure choice, mobility, and demand, and the availability of large databases have sharpened our estimates.

However, most models, and particularly empirical estimates, have remained static in nature. Consumers are viewed at one point in time. If they move during the observation period, they are often considered to be "closer to equilibrium," and their housing demand is estimated jointly with the decision to move. If they do not move, their tenure choice (own or rent) is often estimated jointly with housing demand, under the premise that they are "out of equilibrium," but that this disequilibrium is not severe enough to overcome the sizable transaction costs that moving entails.

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Such analyses are at variance with the anecdotal wisdom that suggests that consumers:

1. have well-defined expectations as to future incomes and price changes,

2. are fairly well informed about the transaction costs of changing housing locations, and

3. face transaction costs for owner housing that are higher than for renter housing.

These distinctions may explain some of the discrepancies between broadly measured housing expenditures in the aggregate, which have often indicated income elasticities close to $+1.0$, and the microeconomic elasticities generally closer to 0.5. The aggregate measurements internalize most of the tenure choice and mobility effects. The microeconometric measures must model the tenure choice and mobility effects explicitly.

This study provides a discrete-time consumer optimization model with transaction costs. It demonstrates comparative statics results from this model, showing how they vary from simpler models, and it links the comparative statics with mobility analysis. It then proposes econometric techniques to estimate the model and presents a database that has been created from the ''American Housing Survey'' (AHS, [26]). The primary finding is that in a multiperiod model the impacts of incomes and price variables from different periods are separable and significant. These impacts indicate the importance of looking at housing stayers when modeling housing demand.

2. MULTIPLE PERIOD FRAMEWORKS

Goodman [10, 11] derives a model in which the transaction costs of changing dwellings are essentially infinite. The two-period framework in those two pieces, while useful for exposition, ignores the decisions on how long to stay and how often to move. Goodman [12] reformulates a key aspect of the earlier analyses that links the static housing demand model to mobility analysis and considers a multiperiod model, showing the equilibrium conditions, demonstrating that they are unique, and presenting comparative statics.

Models of transaction costs in adjusting activity levels are not new. Hu [17] considers the appropriate adjustments to capital stock when the transaction costs are large. In housing analysis, Muth [21] examines moving costs in the context of long-term housing expenditures. Amundsen [2] considers the optimal number of moves when a consumer has perfect foresight and can access perfect capital markets. He shows how the number of moves is related to moving costs, income, and preferences for housing, and he demonstrates, under simplified conditions, that the moves are equally spaced.[2]

[2] Ai, Feinstein, McFadden, and Pollakowski [1], Edin and Englund [6], and Henderson and Ioannides [16] conduct empirical studies treating moving costs.

These models do not address several aspects of housing analysis. The first is consumer choice under imperfect capital markets. The permanent income hypothesis suggests that consumers can easily borrow against future earnings, but ‘‘real life lenders’’ are not so accommodating. The considerable literature on liquidity constrained borrowing suggests major capital market imperfections, particularly early in peoples’ earning lives.

The second aspect involves linkages between demand and mobility. Most models view consumers either as purchasing housing services in equilibrium or as moving when out of equilibrium. How are the two linked, and what indicators can be used to predict mobility?

Third, consumers’ utility functions may change over time, particularly with respect to life cycle variables such as family size, number of children, and retirement. A discrete-time model permitting parameterization of the relative demands for housing and other goods allows a more realistic characterization of the path of housing consumption.

The following Lagrangian equation examines the utility foregone by not moving in response to changed economic conditions:

$$\Omega = U^1(h_1, c_1) + D^{-1}U^2(h_2, c_2) + \lambda_1[y_1 - p_1 h_1 - c_1 - s] + \lambda_2[y_2 - p_2 h_2 - c_2 + (1 + r)s] + \varphi(h_2 - h_1). \tag{1}$$

Here c_1, c_2 are the consumption in Periods 1 and 2, respectively; h_1, h_2 are housing in Periods 1 and 2; y_1, y_2 are income in Periods 1 and 2; s is saving; $D = 1 + \xi$, where ξ is the rate of time preference; and r is the interest rate.

Although (1) can be optimized easily by setting $h_1 = h_2 = h$, the explicit constraint $h_1 = h_2$, along with its multiplier φ, provides a measure of foregone utility due to immobility. Comparing φ to the foregone utility due to a moving cost incurred in Period 2 (reducing Period 2 disposable income, and hence utility) would indicate whether the consumer would stay or would move.

Optimizing over c_1, c_2, h_1, h_2, and s yields the first order condition

$$U_c^1\left[\left(U_h^1/U_c^1\right) - p_1\right] = \varphi = -D^{-1}U_c^2\left[\left(U_h^2/U_c^2\right) - p_2\right].$$

Since U_c^1 is the marginal utility of income at time 1 and $D^{-1}U_c^2$ is the marginal utility of income at time 2,

$$MU_y^1(MRS_1 - p_1) = \varphi = -MU_y^2(MRS_2 - p_2). \tag{2}$$

Equation (2) holds whether saving or borrowing can occur. With perfect capital markets, $U_c^1 = D^{-1}U_c^2(1 + r)$, so

$$(1 + r)(MRS_1 - p_1) = \varphi = -(MRS_2 - p_2). \tag{3}$$

If housing prices or utility functions differ in the two periods, or if discount rate ξ and interest rate r differ, then the immobility constraint is binding and φ represents a cost (in terms of foregone utility) that can be related to moving costs.[3] In equilibrium, consumers will adjust housing purchases such that the difference between the marginal rate of substitution (MRS) and the price ratio p_1, weighted by the marginal utility of income in Period 1, equals the negative of the similar term in Period 2. This difference *equals* the disutility φ imposed by the constraint that the consumer must stay in the same unit. Multiplier φ can be eliminated from Eq. (2) or (3), so that

$$\sum_{t=1}^{t=2} MU_y(MRS_t - p_t) = 0. \tag{3'}$$

The equivalence between (3) and (3′) allows one to examine multiperiod optimizations since it relates the number of moves and the probability of moving to disutility stemming from moving costs.

Equilibrium foregone utility φ^* is easily related to the foregone utility that would occur due to a moving cost m (henceforth expressed in terms of foregone utility through reduction in disposable income) between the two periods. If $\varphi^* > m$, this implies a move each period, trading disposable income reduction m for the greater gain in utility.

Also, φ^* provides a way to look at changes in prices, incomes, or tastes. Figure 1, derived from Eq. (3), supposes that $p_1 > p_2$. Housing amounts h_1^* and h_2^* would be chosen if there was no immobility constraint. The left-hand side of Eq. (3) equals 0 at h_1^* and is decreasing in h. The right-hand side of (3) equals 0 at h_2^* and is increasing in h. The intersection of the two curves (point A) at h^*, with immobility cost φ^*, implies that the Period 1 housing purchase is too large and the Period 2 housing purchase is too small relative to the unconstrained mobility case.

Consider now an increase in p_1 to p_1'. At previous equilibrium level h^*, the left-hand side expression $(1 + r)(MRS_1 - p_1')$ is now less than $-(MRS_2 - p_2)$. In response to the price change, housing quantity h must decrease so that MRS increases each period. At the new equilibrium point B, $(1 + r)(MRS_1 - p_1') = -(MRS_2 - p_2)$, but the new equilibrium is at h^{**}, which is less than h^*. The price increase in Period 1 reduces the quantity demanded in both periods and the immobility cost increases from φ^* to φ^{**} because the absolute difference between p_1 and p_2 has increased. A consumer previously at the margin between moving and staying may now choose to move. One can see that had p_1 decreased (approaching p_2), the $(1 + r)(MRS_1 - p_1')$ curve would shift to the right, increasing the equilibrium housing level. If $p_1' = p_2$, the curves

[3] The sign of φ depends on the sign of its multiplicand, so φ is referenced by absolute value. When $\varphi = 0$ (the constraint is not binding), standard single-period equilibrium conditions hold.

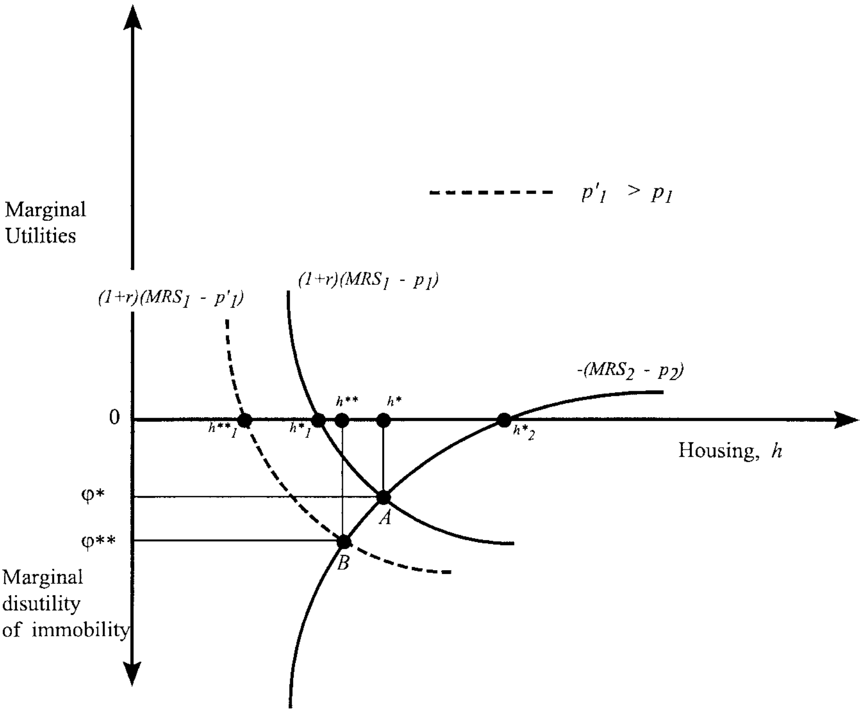

FIG. 1. The immobility constraint under perfect capital markets ($p_1 > p_2$).

intersect at h_2^* with the immobility cost φ now equal to 0, since the immobility constraint has become nonbinding.[4]

Consider instead a consumer optimizing over T periods. The transaction cost of moving each period is m_t. If a consumer, at time 0, plans to move each period, the discounted sum of future utilities is

$$U^* = \sum_{t=1}^{t=T} D^{1-t} U^t(h_t^*, c_t^*) \tag{4}$$

and the budget constraint each period is

$$y_t = p_t h_t + c_t + m_t. \tag{5}$$

Staying in the same unit for two or more periods permits the consumer to save moving costs, while again incurring immobility (in terms of foregone utility)

[4] Goodman and Wassmer [14] solve a two-period model with perfect capital markets and Cobb–Douglas utility function for the utility loss due to the immobility constraint.

costs. Given the multidimensional vector of incomes, prices, and preferences, the consumer solves for:

1 the number of stays (alternatively, the number of moves), k;

2. the length of each stay (alternatively, the number of periods between moves), $S_k = (T_k - T_{k-1})$, with $\sum_k S_k = T$;

3. housing consumed during each stay, $\bar{h}^k$; and

4. nonhousing consumption during each period, c_t.

Goodman [12] demonstrates that the multiperiod equilibrium is summarized by Eq. (3′), with the weighted sum of the differences between the marginal rate of substitution and the price ratio over the multiperiod stay equaling 0. Each period's income and housing price, as well as the prices of other goods and other sociodemographic characteristics, influence the quantity of housing purchased during the *entire stay*, even for households that do not move.

3. ECONOMETRIC MODELING

Testing the theoretical model presents challenges. One would desire to follow a panel of households over time, seeing some move, possibly several times, and some stay. The theoretical model does not explicitly model tenure choice, so any empirical housing work must address issues of owning as opposed to renting, particularly regarding the roles of moving and transaction costs.

The database covers households in the Detroit metropolitan area in 1981, 1985, and 1989. Looking at households that were in the sample in both 1981 and 1985:

1. Were they owners or renters?

2. Did they stay in the dwelling unit between 1985 and 1989?

3. Conditional on Parts 1 and 2, how much housing did they own in 1989?

A five-equation model is used to address these questions. The first two equations establish instruments to be used for permanent income and housing price. The third and fourth equations jointly estimate tenure choice (owner−renter status) and mover−stayer status. The final equation estimates housing demand conditional on staying in the same dwelling unit.

Permanent income estimates follow the cross-sectional method proposed by Goodman and Kawai [13], as return r_h on human capital vector $\mathbf{H}$ and return r_n on nonhuman capital vector $\mathbf{N}$:

$$Y^P = r_h\mathbf{H} + r_n\mathbf{N}. \tag{6}$$

Substituting Eq. (6) into the identity that current income Y equals the sum of its permanent (Y^P) and transitory (Y^T) components, or $Y = Y^P + Y^T$, yields

$$Y = r_h\mathbf{H} + r_n\mathbf{N} + Y^T. \tag{7}$$

Here, the predicted value of the regression on human capital variables, including age, education, gender, and race, and nonhuman capital variables, including financial assets, is taken as permanent income. The residual is treated as transitory income Y^T.

Housing prices are estimated with hedonic price equations following the formula

$$\log V = v_0^o + \sum v_0^o x_k + v_G^o G + u^o \tag{8a}$$

$$\log R = v_0^r + \sum v_0^r x_k + v_G^r G + u^r, \tag{8b}$$

where $V(R)$ = the value (rent) of the dwelling unit, depending on the vector X of housing attributes and the location G. House and rental price indices are calculated over geographic areas G for standardized bundles X^* such that $P_o^G = V(X^*, G)$ and $P_r^G = R(X^*, G)$.

Estimating consumer behavior suggests a joint relationship between housing tenure (own/rent) and the probability of being in the sample four years hence (the move/stay decision). The two are related—Shelton [24] and others since have modeled the economic factors that lead renters to shorter (implicitly more likely to move in any time interval) housing tenures than owners.

A bivariate probit model (Catsiapis and Robinson [4], Ermisch [7], Greene [15], and Maddala [19]) is used to estimate the joint relationship for housing tenure f and the probability of staying g. Variable $f = 1$ if and only if the household owned, with $f = 0$ referring to renter housing. Variable $g = 1$ if and only if the household was in the 1989 sample (i.e., the household "stayed"), with $g = 0$ otherwise.[5] This is analogous to comparing staying costs (in foregone utility) with moving costs.

$$\textit{Owners}:\ f = \mu_Y Y + \mu_P(P_o/P_r) + \mu_v(V/R) + \sum \mu_D D + \mu_l L + \varepsilon_f \tag{9a}$$

$$\textit{Stayers}:\ g = \alpha_Y Y + \sum \alpha_\sigma \sigma + \sum \alpha_D D + \alpha_l L + \varepsilon_g \tag{9b}$$

The correlation of ε_f and ε_g is denoted by ρ.

The specification of tenure choice (9a) follows Goodman [9]. All else being equal, increased income Y and length of stay L are likely to predict owner housing, and the D terms such as household size, age, gender, or race of head

[5] Strictly speaking, f and g are continuous latent variables and the observable dichotomous ones are defined relative to these variables' crossing the zero threshold or not.

of household may reflect tastes.[6] Goodman distinguishes between the owner–renter price ratio P_o/P_r and the value–rent ratio V/R. For comparable dwelling units with attributes X^*, an increase (decrease) in $P_o(X^*)/P_r(X^*)$ is expected to predict renter (owner) status.

In contrast, V/R is derived to reflect expected housing investment returns—high (low) V/R is expected to predict owner (renter) status for specific dwelling units (Goodman [9]). Through a well-specified function, one can reconstruct any renter (owner) unit as if it were owned (rented). Since hedonic coefficients can be interpreted as the sums of replacement costs (Rosen [23]) and quasi-rents (Kain and Quigley [18]), a set of high quasi-rents for a specific bundle suggests a market-indicated expectation for capital gain. Holding the relative prices for standardized units constant, the value–rent ratio compares units for investment potential.

Specification of the stayer equation (9b) follows the theoretical derivation of Eqs. (2) and (3), which indicate that differences over time in explanatory variables such as income and housing price may impose higher staying costs. The σ terms refer to ‘‘spreads’’ of incomes, prices, and value–rent ratios, the variables D referring to sociodemographic variables that may reflect tastes, and L refers to the length of stay in the residence. Since it is postulated that owners are more likely to stay, the simultaneity between housing tenure and probability of staying is estimated in correlation ρ between f and g.

Conditional on ‘‘staying’’ in the 1989 sample,[7] owner and renter housing demand are

$$q^{own} = \left[\sum_{i=1}^{3} \eta_{yi}^{P} Y_i^P + \eta_{pi} P_{oi} + \eta_{vi} (V/R)_i \right] + \sum_k \eta_{kD} D_k + \eta_f \lambda_f + \eta_g \lambda_g + \varepsilon_o, \tag{10a}$$

$$q^{rent} = \left[\sum_{i=1}^{3} \delta_{yi}^{P} Y_i^P + \delta_{pi} P_{ri} \right] + \sum_k \delta_{kD} D_k + \delta_f \lambda_f + \delta_g \lambda_g + \varepsilon_r. \tag{10b}$$

As derived from the theoretical model, multiple measures of income, housing price, and value–rent ratio are included for each of the three years. Variables λ_f and λ_g refer to the selection adjustments derived from Eqs. (9a) and (9b).

Examining λ_f and λ_g further using Greene's notation, let the subscript f refer to owner–renter probability and the subscript g refer to mover–stayer

[6] Length of stay may plausibly be treated as endogenous. Doing so at this time is beyond the scope of this analysis.

[7] When subjects move, they are lost to the 1989 sample with respect to estimating demand.

probability. Then let $h_f = 2f - 1$ and $h_g = 2g - 1$. Thus $h_f = 1$ if $f = 1$ and -1 if $f = 0$; the case is similar for $g = 1$ or 0. Now let

$$z_j = \boldsymbol{\beta}'_j \mathbf{x}_j;\ w_j = h_j z_j,\ j = f, g;\ \text{and}\ \rho^* = h_f h_g \rho.$$

The probabilities of individual cells for stayers are

$$\text{Prob(own, stay)} = \Phi_b\left(w_f | f = 1, w_g | g = 1, \rho^*\right) = \Phi_b(z_f, z_g, \rho)$$

$$\text{Prob(rent, stay)} = \Phi_b\left(w_f | f = 0, w_g | g = 1, \rho^*\right) = \Phi_b(-z_f, z_g, -\rho),$$

where Φ_b refers to the bivariate normal distribution.

Selection adjustments λ_f and λ_g for individual observations are calculated as

$$\lambda_f = \frac{h_f \phi(w_f) \Phi\left[\dfrac{w_g - \rho w_f}{\sqrt{1 - \rho^2}}\right]}{\Phi_b[\cdot]}, \qquad \lambda_g = \frac{h_g \phi(w_g) \Phi\left[\dfrac{w_f - \rho w_g}{\sqrt{1 - \rho^2}}\right]}{\Phi_b[\cdot]}, \tag{11a, 11b}$$

where

$$\Phi_b[\cdot] = \Phi_b\left[\frac{w_g - \rho w_f}{\sqrt{1 - \rho^2}}, \frac{w_f - \rho w_g}{\sqrt{1 - \rho^2}}, \rho^*\right].$$

Since variables such as income are used in several stages of the estimation, it is important to show how they are used to calculate marginal impacts and elasticities. Because relatively few renters in the sample stay in the same unit for more than eight years, the analysis concentrates on the demand of owner−stayers. Expected housing demand is the probability of being an owner−stayer multiplied by the amount of housing demanded by those who are owner−stayers. Following Greene, identify the tenure choice regression as f and the mover−stayer regression as g. Then let the vector $\mathbf{x} = \mathbf{x}_f \cup \mathbf{x}_g$ and let $\boldsymbol{\beta}'_f \mathbf{x}_f = \boldsymbol{\gamma}'_f \mathbf{x}$ and $\boldsymbol{\beta}'_g \mathbf{x}_g = \boldsymbol{\gamma}'_g \mathbf{x}$.[8]

The bivariate probability reflecting owner−stayer status is

$$\text{Prob}[f = 1, g = 1] = \Phi_b[\boldsymbol{\gamma}'_f \mathbf{x}, \boldsymbol{\gamma}'_g \mathbf{x}, \rho], \tag{12}$$

[8] As a result of these transformations, γ'_f contains all the nonzero elements of $\boldsymbol{\beta}'_f$ and possibly some zeros in the positions of variables in $\mathbf{x}$ that appear only in the other equation; $\boldsymbol{\gamma}'_g$ is defined similarly (Greene [15, p. 851]).

and the demand by owner–stayers (Eq. (10a)) is rewritten as

$$q^{own} = \left[\boldsymbol{\eta}' \mathbf{x}_i + \eta_{\lambda_f} \lambda_f + \eta_{\lambda_g} \lambda_g \right]. \tag{13}$$

The conditional elasticity of owner–stayers (following Greene) consists of two components. The direct effect of variable x on the mean of q_i^{own} is η. In addition, a variable such as income Y, which appears in one or more probability equations, will influence q_i^{own} through its presence in λ_f and λ_g.[9] The effect of a 1% income increase on q_i^{own}, for example, is

$$\Delta q_i^{own} = \left[\boldsymbol{\eta}' \mathbf{x}_i^{(Y=1.01Y_0)} + \eta_{\lambda_f} \lambda_f^{(Y=1.01Y_0)} + \eta_{\lambda_g} \lambda_f^{(Y=1.01Y_0)} \right] - \left[\boldsymbol{\eta}' \mathbf{x}_i^{(Y=Y_0)} + \eta_{\lambda_f} \lambda_f^{(Y=Y_0)} + \eta_{\lambda_g} \lambda_f^{(Y=Y_0)} \right]. \tag{14}$$

The derived percentage change in Δq_i^{own} thus represents the income elasticity.

From the theoretical model it follows that one should model a permanent increase in income (for example) as a one-dollar increase in each of the three years. Similar effects, using multiple measures of housing price or value–rent ratio, can be derived for other ‘‘economic’’ variables. Estimating separate significant coefficients for income in different years within the same equation would provide separate effects over time, as predicted by the theoretical model.

A ‘‘naïve’’ model will also be estimated, using a cross-section sample with only single-period variables for housing quantity or expenditures. Such a model is not nested within Eqs. (10a) or (10b) because it uses only a single-period, rather than a multi-period, measure of housing quantity and because it cannot model the mover–stayer selection adjustment, rendering it part of the error term. Nonetheless, it provides a useful benchmark for comparisons.

4. SAMPLE

The American Housing Survey (AHS) was chosen because it provides details on both the dwelling units and the households within them that are not available in other databases.[10] Moreover, the methods developed are replicable on AHS databases for other metropolitan areas.

[9] Greene signs $\partial E[y_i \mid z_i^* > 0]/\partial x_{ik}$ in a conventional probit model, where z^* is the selection parameter and y is the dependent variable conditional on selection. He writes ‘‘it is quite possible that the magnitude, sign, and statistical significance of the [full] effect might all be different from those of the estimate of [the direct effect] $\beta \ldots$’’ [15, pp. 928–929]. In an e-mail to the author, Greene indicated that $\partial E[y_i \mid f_i^*, g_i^* > 0]/\partial x_{ik}$ is exceedingly difficult to sign, but calculating the incremental impacts depends only on the ability to evaluate λ_f and λ_g at incremental levels of the explanatory variables.

[10] This feature contrasts with a database like the Panel Survey on Income Dynamics (PSID) that is explicitly panel, but which contains only limited housing data and even less geographic detail.

It is essential to show how the household database was constructed because the AHS follows dwelling units (each with a unique identification number) rather than individuals. One cannot assume without additional information that the same household is occupying the dwelling unit. Although other studies have utilized the panel nature of the dwelling units, it is unknown whether any have attempted to take advantage of the panel nature of the households.

Due to confidentiality concerns the AHS does not provide geographic identifiers on its publicly available files. As a result, dwelling unit location within a metropolitan area is limited to central city, suburb, or county indicators. For example, all houses in the city of Detroit have identical unit prices for a given year, and identical unit prices four and eight years later.[11]

When the project started, the Detroit MSA surveys were available for 1973, 1977, 1981, 1985, and 1989. Because of survey procedure changes, the study was limited to 1981, 1985, and 1989. Some dwelling units were rotated out of the survey, so the demand analyses used only households from dwelling units included in all three. Thus households outside of Wayne, Macomb, and Oakland Counties (the three counties in 1981) were not used. With no reason to believe that units were systematically rotated out of the sample, there is no reason to assume selection bias.

The fundamental criterion for matching indicated when the household had moved into the unit. Suppose 1981 Household A (HA) had lived there since 1978. Looking at the same house in 1985, if the 1985 household had also lived there since 1978, and matched on age of household head and other consistency criteria, it was assumed that this was HA for both 1981 and 1985, and that HA had been there for 7 years. The process was repeated for the 1989 panel.[12]

If HA is also identified for 1989, then it is indicated as having lived in the dwelling unit for 11 years (since 1978). If, however, in looking at the same dwelling unit for 1989, the current household has been there since 1987, two assumptions are made:

1. HA moved from the dwelling unit in 1987.
2. Household B (HB) moved into the dwelling unit in 1987.

Thus, it is assumed that HA lived in the dwelling unit for 9 years (from 1978 to 1987). Household B enters the sample, having lived in the dwelling unit for 2 years. A total of 3,200 households were identified in this process.[13]

[11] Unit prices do change from period to period. Inclusion of additional metropolitan areas in this estimation process would presumably alleviate lack of price variation.

[12] The entire set of matching criteria and procedures are available from the author upon request.

[13] There is potential for error since HA may have moved out anytime after 1985, and the unit may have been vacant, or occupied by someone else, before HB moved in. This suggests that the spell lengths for those identified as movers, at the time of their moves, may be biased slightly upward.

One of the major premises is that households continue to live in the same unit, consuming roughly the same quantity of housing. Even with ‘‘perfect’’ measurement, housing quantity may change within the same unit due to renovation or depreciation. Over 8 years these magnitudes will be small, although they could compound to larger amounts over longer periods.

Housing quantity is calculated by dividing estimated value by the price of housing estimated from Eqs. (8a) and (8b), yielding (for three years) q_{81}, q_{85}, and q_{89}. Since this process of calculating housing quantity does not constrain q_{81}, q_{85}, and q_{89} to equal each other, the arithmetic average of the three is used as the housing quantity.

The main multiyear demand analyses were based on a sample of 906 households who lived in the same dwelling unit in both 1981 and 1985 (Table 1). At the 1985 benchmark, approximately 19% of sample households were Black, and about 63% of the households were married. Mean age of the household head was 53.4 years, and the mean household length of stay was 17.2 years (varying from a minimum of 5 years to a maximum of 62 years). Summary measures of income and price are discussed below.

5. RESULTS

a. *Income and Price*

For permanent income, separate owner and renter regressions were estimated for each year for *all* households in the AHS database (rather than simply for those who stayed in the same unit) that year. The estimating regression is

$$\begin{aligned} \textit{Wage Income} &= Y - r_n\mathbf{N} \\ &= r_0 + r_1(\textit{AGE}) + r_2(\textit{EDUC}) + r_3(\textit{DEM}) + u, \quad (7') \end{aligned}$$

where $r_n\mathbf{N}$ nets out nonwage returns to nonhuman capital. Parameter r_1 refers to a cubic function of the age of the household head, r_2 to levels of education (high school, some college, college degree, graduate work), and r_3 to demographic variables such as gender, race, marital status, and presence of a second worker. The regressions (available from the author on request) were estimated in nominal (1981, 1985, or 1989) dollars; all results subsequently were deflated to real (1982−1984) dollars.

A criticism of Eq. (7′) is that the error u may contain systematic components attributable to unmeasured skills or effort. These components cannot be identified in cross-sectional regressions, but can be estimated for households for whom there is more than one observation. For households in the sample for two or three years, $\bar{u}_2 = (\hat{u}_{81} + \hat{u}_{85})/2$ or $\bar{u}_3 = (\hat{u}_{81} + \hat{u}_{85} + \hat{u}_{89})/3$ was calculated, as appropriate. Systematic effects $\bar{u}_2$ or $\bar{u}_3$ were then added to the fitted values $\hat{u}$ of Eq. (7′) for each year as permanent income and subtracted from $\hat{u}$ as transitory income.

TABLE 1
Variables for Multiyear Models

Variable	Mean	Std. Dev.
Current income		
1981	24983.83	14826.49
1985	32486.69	25822.05
1989*	37615.77	29064.14
Permanent income		
1981	24000.89	16656.31
1985	31047.53	21267.96
1989*	32666.43	22402.58
Owner price		
1981	39044.60	6890.77
1985	55864.17	13863.31
1989*	77010.54	23175.03
Renter price		
1981	271.17	28.50
1985	446.35	30.20
1989*	543.08	52.53
Value-rent		
1981	106.98	26.23
1985	158.83	45.28
1989*	136.82	52.40
Po/Pr		
1981	142.96	13.57
1985	123.86	25.85
1989*	138.97	33.40
Other variables		
AGE85	53.39	15.46
BLACK85	0.19	0.39
MAR85	0.63	0.48
TEN85	0.84	0.36
HOWLON85	17.20	10.47
PCHYP	0.51	1.79
PCHVR	0.40	0.20
PCHP	0.14	0.12

* Variable mean for housing stayers only.
Note. All dollar values are in real $1982-4.

Returning to Table 1, mean sample current income rose from $24,984 in 1981 to $37,616 in 1989. Permanent income rose similarly, from $24,001 in 1981 to $32,666 in 1989. Households who were in the sample for all three periods had mean annual real transitory income of $995, approximately 3 to 4% of total income.

Housing prices and value−rent ratios were derived from hedonic price regressions estimated in semi-log form (Thibodeau [25]). Separate regressions were estimated by year and for differing tenures, but geographic submarkets were modeled solely with binary variables. The 1980s saw considerable population loss in the City of Detroit relative to the rest of the metropolitan area, and this is indicated by steep house price discounts (-33.6%, -46.5%, and -56.6%) for the three years.

Renter hedonic price regressions were also estimated for 1981, 1985, and 1989. Detroit rents were not as steeply discounted, although they were 19.7%, 10.9%, and 21.2% less than in surrounding areas in the three years, respectively.

The price indices used the arithmetic mean of owner and renter bundles as X^*. Indices P_o and P_r apply the Duan [5] ''smearing'' factor $s = \sum \exp(\hat{e}_i)/n$, where $\hat{e}_i = y_i - x_i \hat{\beta}$ refers to estimated residuals, to retransform semi-log estimates from Eqs. (8a) and (8b):

$$P_o^G = s_o \exp\Big(v_0^o + \sum v_k^o x_k^* + v_G^o G\Big), P_r^G = s_r \exp\Big(v_0^r + \sum v_k^r x_k^* + v_G^r G\Big).$$

The prices (in \$1982–1984) for identical units for the four areas (Detroit, Wayne County outside Detroit, Macomb County, and Oakland County) are provided in the Appendix. Detroit prices started the decade depressed relative to the suburbs and remained so throughout the decade.

Value–rent ratios for individual units are created by statistically matching owner units with renter units with the same characteristics using Eqs. (8a) and (8b). Because the vectors of coefficients were allowed to vary by dwelling unit, there was considerably more variation in value–rent ratios than in housing prices.

Returning to Table 1, owner house prices approximately doubled from 1981 to 1989, as did the renter unit prices. For the multiyear sample, the owner–renter price ratio decreased from 143 in 1981 to 124 in 1985. The value–rent ratio, in contrast, rose from 107 in 1981 to 159 in 1985.

b. *Naïve Demand Regressions*

It is useful to describe the data with naïve demand regressions, specifying the regressions as if they were cross-sectional demand regressions. Because the multiyear models concentrate on owner housing, this section concentrates on owner regressions. A linear functional form is used, with no selection adjustment for owner vs renter housing. In Table 2, columns (1), (3), and (5) refer to current income regressions for 1981, 1985, and 1989, respectively. Columns (2), (4), and (6) refer to permanent (and transitory) income regressions for the three years.[14]

[14] The regressions presented here have varying numbers of observations since they included households that were in the sample for only one period.

TABLE 2
Naïve Regressions

	QO81	QO81	QO85	QO85	QO89	QO89
INTERCEP	**−0.21786**	**−0.35806**	**0.29044**	**0.22110**	**0.51934**	**0.42041**
	0.25181	0.24361	0.08928	0.09083	0.08612	0.08838
Y	**0.02353**		**0.00735**		**0.00587**	
	0.00204		0.00044		0.00042	
YP		**0.03160**		**0.00955**		**0.00934**
		0.00226		0.00066		0.00070
YT		**0.01503**		**0.00653**		**0.00539**
		0.00206		0.00059		0.00065
PO	**0.02446**	**0.01984**	**−0.00478**	**−0.00478**	**−0.00484**	**−0.00471**
	0.00552	0.00536	0.00145	0.00144	0.00088	0.00088
VR	**−0.00200**	**−0.00125**	**0.00371**	**0.00368**	**0.00502**	**0.00492**
	0.00137	0.00132	0.00040	0.00040	0.00031	0.00031
AGE	**0.00336**	**0.00456**	**−0.00081**	**−0.00017**	**−0.00448**	**−0.00362**
	0.00191	0.00185	0.00080	0.00082	0.00084	0.00086
BLACK	**−0.12458**	**−0.13786**	**−0.00859**	**0.00097**	**0.01035**	**0.02060**
	0.08974	0.08657	0.04117	0.04105	0.04468	0.04446
MAR	**0.04308**	**0.00997**	**0.05504**	**0.03028**	**0.06101**	**0.02437**
	0.06595	0.06376	0.02497	0.02575	0.02617	0.02726
N	718	718	1243	1243	1524	1524
R^2	0.2668	0.3189	0.3299	0.3371	0.3367	0.3454
Adj. R^2	0.2606	0.3122	0.3267	0.3334	0.3340	0.3423
Ey	0.4591	0.6165	0.2893	0.3758	0.2404	0.3827
Ep	0.7015	0.5691	−0.3081	−0.3079	−0.4391	−0.4270
Ev	−0.1498	−0.0938	0.6849	0.6801	0.8162	0.8009
Ep:v	0.5516	0.4754	0.3768	0.3721	0.3771	0.3739

Note. Standard errors are given below the coefficients.

In Table 2, estimated current income elasticities vary from 0.240 (in 1989) to 0.459 (in 1981); permanent income elasticities are higher (as high as 0.616 for 1981). Decomposing current income into permanent and transitory income improves the regressions’ explanatory power. Transitory income has from one-half to two-thirds of the impact of permanent income.

Housing price is expected to have a negative impact on quantity demanded, and the value−rent ratio is expected to have a positive impact. These predicted effects occur for 1985 (price elasticity of approximately −0.31) and 1989 (price elasticity of −0.43 to −0.44), but not for 1981. The combined impacts of 1% changes in housing prices and value−rent ratios yield housing demand increases varying from 0.372 to 0.552%.[15]

[15] The reported joint impacts of price and value−rent ratio with respect to housing demand and ownership, here and elsewhere in this article, are consistent with Goodman’s [9] findings.

The sociodemographic variables are age and race of household head and marital status. Increased age is positively correlated with housing demand in 1981, but negatively correlated in 1989 (the 1985 coefficient is negative but insignificant). Blacks purchased less owner housing in 1981 and 1985 than did whites, but the coefficients are not statistically significant: they purchased more in 1989, but those coefficients too are insignificant. Married households purchased more housing than did unmarried households in all three years, but when permanent income was taken into account this difference, too, is insignificant. In sum, the results provide cross-sectional price and income elasticities that are consistent with estimates from the housing demand literature.

c. *Multiyear Models*

Tables 3 and 4 estimate the multiyear model. The analysis investigates households in the sample in 1981 and 1985, asking whether the household: (1) stays in 1989; (2) owns or rents in 1989; and, conditional on staying and on owning or renting, (3) how much the household demands.

Table 3 presents the bivariate probit analyses of tenure choice and mover−stayer status.[16] Columns 1 and 2 use the current income in both the tenure choice and the mover−stayer relationships. Columns 3 and 4 decompose current income into its permanent and transitory parts for each relationship. The decomposition of current income into permanent and transitory parts yields a statistically significant improvement in explanatory power by likelihood ratio criteria.[17] Both estimates are discussed, and both are used in the subsequent demand estimation stage.

The validity of the joint estimation of tenure-choice and mover−stayer status is supported by the significant disturbance correlation (0.413 for current income; 0.430 for permanent income), indicating that ownership was correlated with staying and renting with moving. In the current income version, the unconditional mean probability of owning, given staying, was 0.898; the unconditional mean probability of staying, given owning, was 0.748. The joint probability of staying in owner housing, using Eq. (12) to account for the correlation of the error terms, was 0.700.

Increased 1981 and 1985 incomes, included in either current (columns 1 and 2) or permanent (columns 3 and 4) form, are separately correlated with owner rather than renter housing. Transitory income has a significant impact in 1981, but an insignificant impact in 1985. Married households are more likely to own. Black households are also more likely to own, although the coefficients are not

[16] The bivariate probit analysis was done with LIMDEP Version 7.0, and the calculation of the selection variables λ_f and λ_g was performed through a subroutine available from LIMDEP.

[17] The current income estimates impose four coefficient restrictions, compared to the permanent (and transitory) income estimates. Twice the difference in log-likelihood functions $\sim \chi^2(4)$. The test statistic of 9.998 is larger than the 95% critical value of 9.488.

TABLE 3
Bivariate Probit Selection Regressions

	Current income			Permanent income	
	(1)	(2)		(3)	(4)
	Tenure	Move−Stay		Tenure	Move−Stay
Variable	(Own = 1)	(Surv = 1)		(Own = 1)	(Surv = 1)
Constant	**−2.08674**	**−0.34388**	Constant	**−2.29833**	**−0.33662**
	0.51970	0.24097		0.55499	0.24769
Y85A	**0.01411**	**−0.00093**	YP85A	**0.02482**	**−0.00400**
	0.00313	0.00249		0.00461	0.00464
Y81A	**0.02911**	**0.00704**	YP81A	**0.01836**	**0.01132**
	0.00630	0.00417		0.00728	0.00545
			YTN85	**−0.00109**	**−0.00752**
				0.00784	0.00504
			YTN81	**0.02107**	**−0.00093**
				0.01032	0.00561
AGE85	**−0.01001**	**0.00137**	AGE85	**−0.00747**	**0.00104**
	0.00468	0.00362		0.00491	0.00372
BLACK85	**0.20524**	**0.32795**	BLACK85	**0.26331**	**0.30532**
	0.18134	0.15309		0.18887	0.15650
MAR85	**0.60319**	**0.16897**	MAR85	**0.52920**	**0.20077**
	0.15868	0.11966		0.16272	0.12479
P85	**0.01296**		P85	**0.01405**	
	0.00442			0.00458	
VR85	**−0.00405**		VR85	**−0.00462**	
	0.00250			0.00258	
HOWLON85	**0.09742**	**0.02212**	HOWLON85	**0.09689**	**0.02196**
	0.00992	0.00556		0.01020	0.00561
PCHYP		**0.00530**			**0.01038**
		0.01639			0.01647
PCHVR		**0.68766**			**0.65966**
		0.28660			0.28843
PCHP		**0.28031**			**0.32485**
		0.53886			0.54567
Disturbance correlation					
ρ		**0.41329**			**0.43009**
		0.08393			0.09026
N		865			865
LL-Ratio		−661.9943			−656.9951

Note. Standard errors are given below the coefficients.

significant. The impacts of owner−renter price ratios and value−rent ratios are reversed from the theoretical predictions, with an increase in the owner−renter price ratio (P_o/P_r) implying homeownership and an increase in the value−rent ratio (V/R) implying renter status. This reversal likely occurs because of the limited variation in the ratios (there are only four values of the price ratio for

TABLE 4

Multiyear Owner Demand Regressions

Current income bivariate probit				Permanent income bivariate probit			
Current income		Permanent income		Current income		Permanent income	
QOWN Variable	(1) Coefficient	QOWN Variable	(2) Coefficient	QOWN Variable	(3) Coefficient	QOWN Variable	(4) Coefficient
Constant	**−0.12282**	Constant	**−0.11210**	Constant	**0.10017**	Constant	**0.06684**
	0.16453		0.16189		0.16113		0.16253
Y81A	**0.00536****	YP81A	**0.00534****	Y81A	**0.00496****	YP81A	**0.00843****
	0.00156		0.00147		0.00156		0.00168
Y85A	**0.00500****	YP85A	**0.00355****	Y85A	**0.00939****	YP85A	**0.00480****
	0.00133		0.00138		0.00102		0.00177
Y89A	**0.00248****	YP89A	**0.00366****	Y89A	**0.00204****	YP89A	**0.00376****
	0.00072		0.00109		0.00075		0.00113
POAVE	**−0.00791****	POAVE	**−0.00874****	POAVE	**−0.00232**	POAVE	**−0.00445***
	0.00242		0.00229		0.00218		0.00228
VR81	**−0.00382****	VR81	**−0.00441****	VR81	**0.00021**	VR81	**−0.00129**
	0.00133		0.00118		0.00105		0.00108
VR85	**0.00590****	VR85	**0.00641****	VR85	**0.00238****	VR85	**0.00351****
	0.00106		0.00095		0.00079		0.00086
VR89	**0.00034***	VR89	**0.00033***	VR89	**0.00043****	VR89	**0.00044****
	0.00018		0.00018		0.00018		0.00019
AGE89	**0.00242**	AGE89	**0.00237**	AGE89	**0.00087**	AGE89	**0.00105**
	0.00148		0.00149		0.00148		0.00152
BLACK89	**0.03345**	BLACK89	**0.04550**	BLACK89	**−0.08302**	BLACK89	**−0.05832**
	0.06463		0.06329		0.06107		0.06175
MAR89	**0.04514**	MAR89	**0.05211**	MAR89	**0.02090**	MAR89	**0.03945**
	0.03758		0.03843		0.03758		0.03938
LAM_TEN	**0.36296****	LAM_TEN	**0.34052****	LAM_TEN	**0.30849****	LAM_TEN	**0.32145****
	0.09369		0.09061		0.08889		0.08828
LAM_STAY	**0.59534****	LAM_STAY	**0.68912****	LAM_STAY	**−0.09997**	LAM_STAY	**0.12108**
	0.16163		0.12661		0.07302		0.08417
N	535		535		535		535
R^2	0.47449		0.47327		0.45983		0.44404
Adj. R^2	0.46138		0.46013		0.44635		0.43017

Note. Standard errors are given below the coefficients.
** Significant at the 5% level.
* Significant at the 10% level.

each year). The joint effect of the two variables (1% increases in both P_o/P_r and V/R), obtained by adding the coefficients, implies a tendency toward homeownership.

Most previous work, including the author's own, has found age to be positively related to ownership. Here, however, controlling for length of stay in the residence, older residents are more likely to rent. This seemingly unusual finding can be interpreted by adding two coefficients. For the current income relationship, an additional year of age has the impact of -0.010; spending that year in the same residence has the impact of 0.097. The net impact is 0.087, a positive impact implying ownership. Older households with shorter stays in the unit are more likely to rent.

The second part of the bivariate probit analysis examines the determinants of ‘‘staying.’’ In addition to incomes, prices, and sociodemographic variables such as age, race, gender, and marital status, the theoretical model implied that changes in the ‘‘economic’’ variables were likely to increase the costs (in terms of foregone utility) of staying, holding moving costs constant. After several specifications, percent absolute change for income, price ratio (owner−renter), and value−rent ratio were chosen such that for variable z, PCH $z = |z_{85} - z_{81}|/z_{81}$. Mean values are given in Table 1.

The resulting estimates provide mixed support for the model as specified. Absolute income levels had little impact on the mover−stayer status (positive coefficients implying stayers) in either estimate. Older households were somewhat more likely to stay, as were married households, although not significantly so. Black households were significantly more likely to stay than whites. Length of stay as of 1985 had a positive impact on staying, indicating that whatever had satisfied the household thus far generally continued to satisfy them (such that they stayed).[18]

Percentage change in income implied increased likelihood of moving, although the impact was not significant. Changed owner−renter and value−rent ratios both had positive impacts on staying, although the theory would suggest that they would be correlated with moving. The impacts of these variables may reflect the Detroit area in the 1980s. There was considerably more price appreciation in the suburbs than in the city. To the extent that suburban households were happy where they lived, they chose to stay there.

Table 4 displays linear owner demand regressions.[19] Columns 1 and 2 use the current income bivariate probit as the selection equation; columns 3 and 4 use the permanent income bivariate probit for selection. The theory presented earlier suggests separate impacts for 1981, 1985, and 1989 income levels, and the cases are similar for prices and value−rent ratios. Preliminary work using three separate house price indices for 1981, 1985, and 1989 indicated that owner prices were almost perfectly correlated across the three years, due to the lack of price variation within jurisdictions. Moreover, due to the multicollinearity there were almost no changes in relative city−suburb prices over the three years. To address this multicollinearity problem, the owner price term was collapsed into a single three-year average term.

[18] The AHS contains subjective evaluations of house and neighborhood quality that might plausibly be related to the decision to move (less satisfied residents may move). These proved statistically insignificant when included in the ‘‘mover−stayer’’ equation.

[19] Renter regressions were also estimated, but the sample size of 60 did not provide satisfactory results, given the need to use 12 regressors. The small sample size is unsurprising since it required that renters spend at least 9 years in the same dwelling unit.

Concentrating first on columns 1 and 2, R^2 was 0.474 for the current income regression and 0.473 for the permanent income regression. Tenure choice (LAM_TEN) and mover–stayer (LAM_STAY) ratio are jointly significant and positively related to demand in both regressions. The joint significance indicates conditioning on sample selection, and the possibility of biased coefficients if selection is ignored.[20]

The three income coefficients are statistically significant at the 5% levels, even though they, too, exhibited considerable collinearity. The three-year average price term is also statistically significant, as are two of the three value–rent ratios.

Since sociodemographic variables were either identical for all three years or explicitly correlated (i.e., age), only the 1989 value was used. Older households purchased slightly more housing than did younger households. Black households purchased 3 to 4% more housing and married households 4 to 5% more. None of these variables are statistically significant.

Columns 3 and 4 use the permanent income probit equations for selection. In contrast to the current income probits, the mover–stayer LAM_STAY is now statistically insignificant, although tenure-choice LAM_TEN remains statistically significant. The R^2 measures are slightly smaller than columns 1 and 2 and the income coefficients are slightly larger. The price coefficients are smaller and are not statistically significant at the 5% level (although the column 4 price term is significant at the 10% level). The sociodemographic variables are still insignificant, although Black households now purchase between about 5 and 8% less housing than do white households.

Table 5 presents income and price elasticities. The income variables are more satisfactory, possibly because they are more accurately measured. Recall that the income elasticity

$$\Delta q_i^{own} = \left[\boldsymbol{\eta}' \mathbf{x}_i^{(Y=1.01Y_0)} + \eta_{\lambda_f} \lambda_f^{(Y=1.01Y_0)} + \eta_{\lambda_g} \lambda_f^{(Y=1.01Y_0)} \right] - \left[\boldsymbol{\eta}' \mathbf{x}_i^{(Y=Y_0)} + \eta_{\lambda_f} \lambda_f^{(Y=Y_0)} + \eta_{\lambda_g} \lambda_f^{(Y=Y_0)} \right] \quad (14)$$

implies that a 1% increase in income affects the quantity for those who own, both through income coefficients and through their indirect impacts on the selection parameters λ_f and λ_g.

Evaluating 1% income increases each year, for column 1, the direct 1981, 1985, and 1989 income elasticities are 0.1483, 0.1795, and 0.1005, summing to

[20] These findings are similar to those of Ermisch [7]. He found that, given observed attributes, owner-occupiers are more likely to be stayers and that households with unobserved attributes that make them more likely to move (stay) have lower (higher) housing demand, while households with unobserved traits that increase their probability of owning have higher housing demand.

TABLE 5
Income and Price Elasticities

	Current income bivariate probit		Permanent income bivariate probit	
	Current Y (1)	Permanent Y (2)	Current Y (3)	Permanent Y (4)
Elas._income	0.3964	0.4013	0.4276	0.4457
Direct y81	0.1483	0.1393	0.1325	0.2141
Direct y85	0.1795	0.1208	0.3256	0.1591
Direct y89	0.1005	0.1276	0.0797	0.1277
Indirect	−0.0319	0.0135	−0.1102	−0.0552
Elas._price	−0.4553	−0.4908	−0.1289	−0.2435
Elas._value-rent	0.6040	0.6083	0.4437	0.4656
Elas._price:value-rent	0.1488	0.1175	0.3148	0.2221

0.4283. The indirect adjustment related to owner−stayer status is −0.0319, leading to a conditional elasticity of 0.3964. The column 2 current income elasticity is slightly higher at 0.4013. The permanent income elasticities in columns 3 and 4 are slightly higher still at 0.4276 and 0.4457 respectively.

The price elasticities are calculated with the single price term. For the demand estimates conditional on current income bivariate probit selection, the owner demand price elasticities are −0.4553 (current income, column 1) and −0.4908 (permanent income, column 2) respectively. For the demand estimates conditional on permanent income bivariate probit selection, the price elasticities are smaller in absolute terms, −0.1289 and −0.2435 for the current income (column 3) and permanent income (column 4) demand formulations, respectively. The positive value−rent ratio elasticity is consistent with its interpretation as a measure of user cost of capital (the higher the value−rent ratio, the lower the user cost, and hence the higher quantity demanded). The joint effect of 1% increases in house price and value−rent ratio is positive.

What is most important in the context of the multiperiod model is that the individual year impacts are generally separable and significant. Moreover, the formulation provides substantive support for the premise that it is appropriate and useful to look at the housing stayers.

6. CONCLUSIONS

This article presents a framework in which the explicit mover−stayer decision is modeled as an equilibrium decision. An ''old-style'' database has been ''rebottled'' to provide a panel of households for which there are very good household data and unmatched housing data.

The results indicate that income and value−rent measures in different years have separable and significant impacts on housing demand. The conditional income elasticities provide values between 0.40 and 0.45. Price impacts on demand are less helpful, in part because of difficulties in measuring housing price using the AHS in a single metropolitan area, even over a period of eight years. However, the three-year average price elasticities are plausible.

From a policy perspective the separable income impacts help to interpret key features of demand-side programs such as housing vouchers that have been proposed to address problems of adequate housing for the poor. The Experimental Housing Allowance Program (EHAP) and successor programs were predicated on time-series income elasticities of Muth [20], Reid [22], and others finding income elasticities greater than +1.0 and as high as +2.0. As Goodman [10] notes, repeated analyses using individual data have generally found income elasticities to be less than +1.0 and most of the analyses from the EHAP project found them to be closer to 0 than to +1.0. The general appraisal was that the EHAP experiments were too short in duration, that the income subsidies were not necessarily viewed as permanent, and that moving costs might constrain adjustment.[21]

Although the demand estimates presented here pertain to owner demand, the findings suggest some validity to the criticism that single-year income measures tend to underestimate responsiveness to income changes. Moreover, they suggest that temporary income changes (such as one-year income increases in the model estimated) can be expected to have only modest impacts on housing demand—programs must promise permanent changes. In short, income subsidies and/or vouchers must be expected and they must be long-term.

It would be useful to add another year and several other metropolitan areas to the database. Even with the three observations (1981, 1985, 1989), the absolute deviation spreads in income, housing price, or value−rent for stayers are limited to differences in two years. Another year (hence another interval), most likely after 1989, would add information on the variable ‘‘spread.’’ Similarly, adding metropolitan areas might provide additional variation in prices across metropolitan areas as well as within them.

It will also be important to take advantage of the current database and to look more explicitly at those who entered the sample after 1981 (the first year), as well as those who were in the sample in 1981, but moved before 1985. Examining the duration of peoples’ stays, as well as the housing demand therein, with mixed estimation models would seem to be yet another fruitful extension of the database and econometric modeling techniques.

[21] Two excellent summary volumes of reports and evaluations regarding EHAP are Bradbury and Downs [3] and Friedman and Weinberg [8].

APPENDIX

TABLE A1

Price Indices for Owner and Renter Housing

	Dollars (1982−1984)			Indexes (Detroit 1981 = 100)		
	1981	1985	1989	1981	1985	1989
			Owner			
Detroit	30547	31992	33048	100.0	104.7	108.2
Wayne	44280	58680	70046	145.0	192.1	229.3
Oakland	48775	62603	75871	159.7	204.9	248.4
Macomb	46040	59997	72026	150.7	196.4	235.8
			Renter			
Detroit	250	380	380	100.0	152.0	152.0
Wayne	283	494	522	113.4	197.9	209.1
Oakland	327	421	448	131.0	168.6	179.4
Macomb	311	461	481	124.5	184.6	192.5

TABLE A2

Variable List

Qualitative Choice	
YP_A	Permanent income in $ (for a given year)
YT_	Transitory income in $
Y_A	Current income in $
AGE_	Age of household head
BLACK_	1 if black; 0 otherwise
MAR_	1 if married; 0 otherwise
P_	Owner−renter price ratio
VR_	Value-rent ratio
PCHYP	Percent change in permanent income
PCHVR	Percent change in value-rent
PCHP	Percent change in owner-renter ratio
HOWLON85	Length of residence
Demand—Additional parameters	
POAVE	Owner price *1,000
LAM_TEN	Selection parameter from tenure choice probit
LAM_STAY	Selection parameter from selection (mover-nonmover) probit

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