MICROSIMULATION OF HOUSING POLICIES WITH A HIERARCHICAL CHOICE MODEL OF HOUSING DEMAND

by

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ABSTRACT

This paper examines the application of a nested multinomial logit model to the demand for housing, specifically for the purpose of projection and policy simulation.

The model has the following features:
(1) Housing consumption is viewed as a choice among a finite number of discrete housing alternatives,
(2) household formation, tenure choice, and choice of the dwelling-size are jointly determined in a hierarchical choice procedure, and
(3) the explanatory variables include demographic and financial variables; the sample is stratified according to the former group of variables, and the latter group is used as regressors.

The paper provides baseline estimates for four representative strata in three SMSAs. In spite of a very parsimonious specification, the model achieves high predictive power.

The estimates are used to simulate three changes of the current tax and subsidy system for housing: a housing allowance program, a cut in the local property tax rate, and a less progressive federal income tax.
The focus of these simulations is on the response of headship rates to these changes and their incidence on the different population strata and on the federal and local level of jurisdiction.

The major results of the paper are as follows:
first, we find a rate of household formation, which is highly price responsive. In particular, the choice probability to "double up" decreases sharply under a housing allowance program. This shows that estimates with exogenous headship rates seriously underestimate the cost and effects of housing allowance programs.
Secondly, all three changes cause spill-over effects between the federal and the local level of government of a substantial magnitude. These spill-over effects are induced by moves from owning to renting and the corresponding changes in local property taxes and deductions from the federal income tax.

Introduction

For a large number of policy purposes, one wants to concentrate less on an aggregate quantitative measure of housing expenditures, and more on the distribution of housing consumption into qualitatively different categories. The most popular example is the choice between renting and owning, and the response of tenure choice to federal income tax treatment. (See Laidler (1969), Rosen (1979), Rosen and Rosen (1980), Henderson and Ioannides (1983).)

This paper goes one step further: not only the choice of tenure, but also the choice of the dwelling size will be affected by taxes and subsidies. Furthermore, the decision whether to form an autonomous household at all may be dependent on relative prices and income. Thus, housing consumption decisions can be divided into three different kinds of decisions, i. e., concerning headship, tenure, and size (as a crude measure of quality).

Lee and Trost (1978) and subsequently King (1980) argue that the tenure choice and the choice of size and quality level are made simultaneously. Boersch-Supan and Pitkin (1982) found evidence that the headship choice is also influencing, and is influenced by the other two decisions, so that all three choices are made in a joint decision process. The specification of this joint decision process including its explanatory demographic and financial variables is discussed in detail in the second section of this paper.

Estimating this joint decision process poses a number of econometric problems: the choice set is fairly large and, due to the threefold nature of the decision process, consists of heterogeneous alternatives, e.g., non-headship versus owning a large house. The first problem restricts the possible specifications of the functional form of the relation between the choice probabilities and the explanatory variables to the class of generalized extreme-value functions, and the second problem prohibits the use of simplifying assumptions like the Independence of Irrelevant Alternatives. This led us to the specification of a nested multinomial logit model (NMNL). The third section of the paper gives a short survey over the microeconomic foundations and the econometrics of a NMNL-model, and an appendix develops a possible reconciliation of these models for the case of dissimilarity parameters larger than one.

A further econometric problem is the handling of the two classes of explanatory variables that are relevant for housing decisions: demographic and financial variables. We follow the same approach as in Boersch-Supan and Pitkin (1982), suggested by de Leeuw (1971): demographic variables are used to stratify the sample, and only financial variables enter the regression directly. This approach is equivalent to the use of dummy variables for each of the strata which interact with all regressors to accommodate the unknown nonlinear fashion (see Li, 1977), in which the demographic variables enter the equations for the choice probabilities.

The model is estimated for four representative population strata:

(1) Young unmarried male and female without children, aged 20-34
(2) Married couples with one or two children, aged 35-59
(3) Elderly married couples without children, aged 60 and above
(4) Widowed, divorced, and separated women without children, aged 60 and above. and for the three Standard Metropolitain Statistical Areas of Albany-Schenectady-Troy, New York; Dallas, Texas; and Sacramento, California, representing the Northeast, the Sunbelt, and the West Coast. The estimates are based on the Annual Housing Survey SMSA cross-sections in 1976 and 1977. Section 4 discusses the baseline results.

The next three sections contain simulations of changes in the tax and subsidy structure for housing consumption. In the first casethe model simulates the impacts of a simple housing allowance program along the lines of the housing gap formula applied in the Experimental Housing Allowance Program in 1973-79. Second, the local property tax rate is assumed to be a half of its actual rate in 1976-77. Finally the model analyses the effect of reducing the highest marginal tax rate of the federal income tax from 70 percent to 50 percent. For all three changes, we calculate the resulting distribution of the population among the different housing alternatives and study the actual moves that take place in response to these changes. Furthermore, the cost of the subsidies at the local and federal levels are evaluated, focusing on the interjurisdictional spill-over effects and on the response of headship rates to price changes induced by the simulations.

The paper concludes with caveats, in particular about the steady-state assumptions of the cross-sectional model, and assesses the predictive abilities of hierarchical choice models and their responsiveness to the various policies.

Specification of the Model

In spite of a stable (and in Europe, a declining) population, we nevertheless observe an increase in housing consumption both in terms of housing units and in housing expenditure. This increase can be attributed to two factors: existing households demand larger (and better) units, and the number of households itself has risen due to an increase in the household headship rate, in particular among young people. Both mechanisms are likely to be influenced not only by income but also by the relative prices of housing. But if the rate of household formation is endogenous to the housing market through price responsive headship rates, estimations and projections on a household basis alone will yield biased results, and just applying exogenous headship rates to forecast models will ignore repercussions and feedbacks.

Pitkin (1980) proposes to split up existing households, such that all potential housing demanders are in fact autonomous decision units, and suggests the introduction of a further category of housing consumption: not heading a household, but finding shelter in an existing household. We will call these smaller decision units "nuclei."

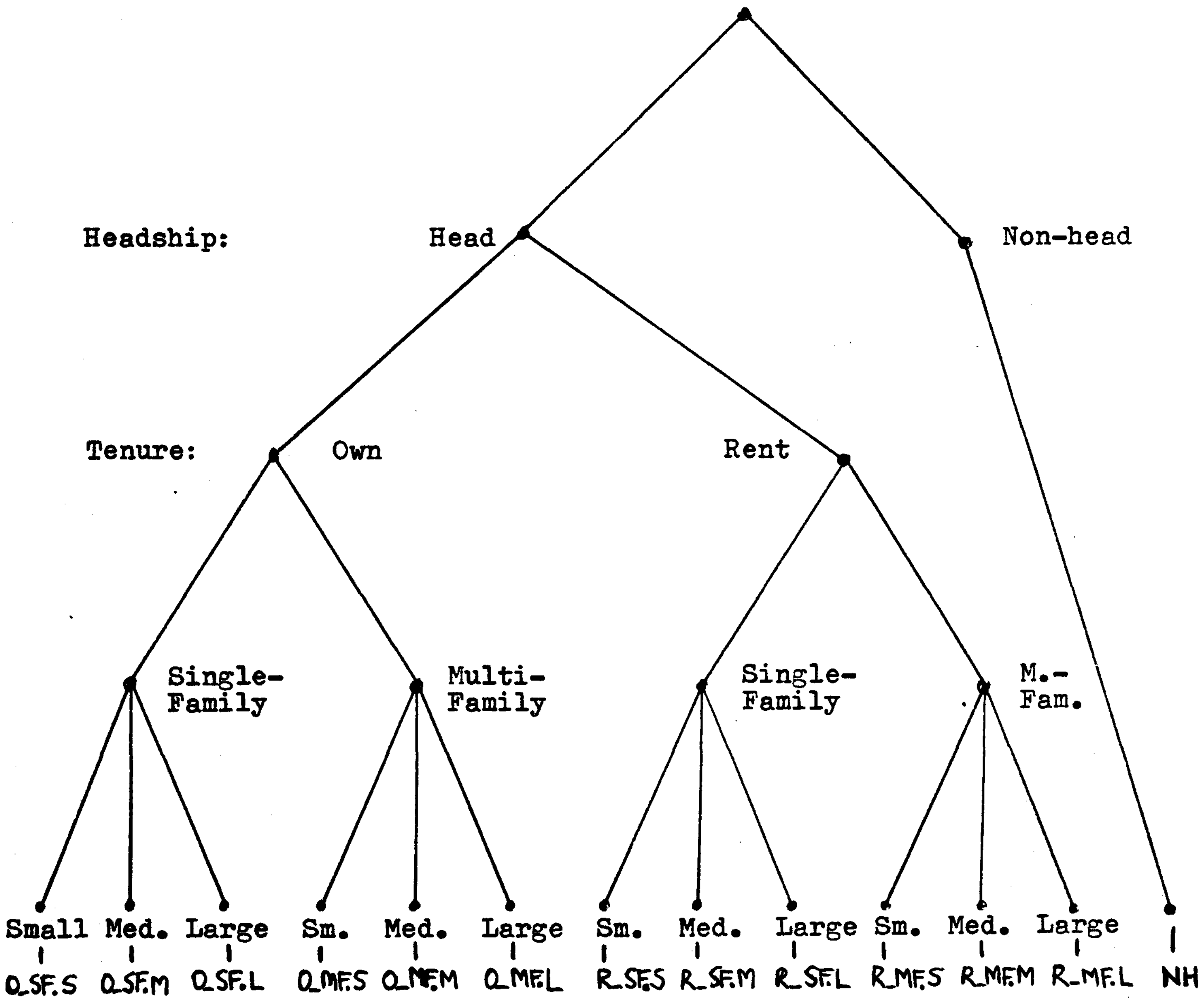
A nucleus consists of a married couple or a single individual together with all its own children below a certain age (say, 18 years). Children above this threshold are considered grown-up and, as potential household heads, form a new nucleus, even if they (still) live in their parents household. Similarly, households that consist of several adults are split up into several nuclei, both when the members are related or unrelated to each other. Examples are elderly parents in the household of their children, or roommates. Accordingly, there are five types of households:

(1) households consisting just of one nucleus,
(2) parents with their adult children,
(3) households composed of nuclei with family-relations,
(4) households composed of nuclei without family-relations,
(5) complex households, i.e., combinations of the latter three types.

Pitkin (1980) presents a variety of behavioral hypotheses for these five types of households, and provides a descriptive analysis of trends in household composition. Pitkin and Masnick (1980) use the nucleus approach for housing projections in the United States. Finally, Boersch-Supan and Pitkin (1982) show the statistical significance of the interdependence between household formation behavior and housing demand. This paper also discusses the procedure of generating a nucleus-based data set from the Annual Housing Survey.

Each nucleus chooses whether to head a household or shelter in an existing household, if one chooses to head a household, thenthe decisions are whether to rent or own a dwelling, and what quality and size the dwelling should be. As a simple measure of quality and size, we use the number of rooms and the type of the building. A household chooses among three size categories and between single-family detached houses and multi-family houses, in particular apartment buildings. We can arrange the choices in form of a "decision tree" as depicted in Figure 1. "Small" refers to dwellings with up to four rooms, "medium"

Figure 1: Basic Decision Tree

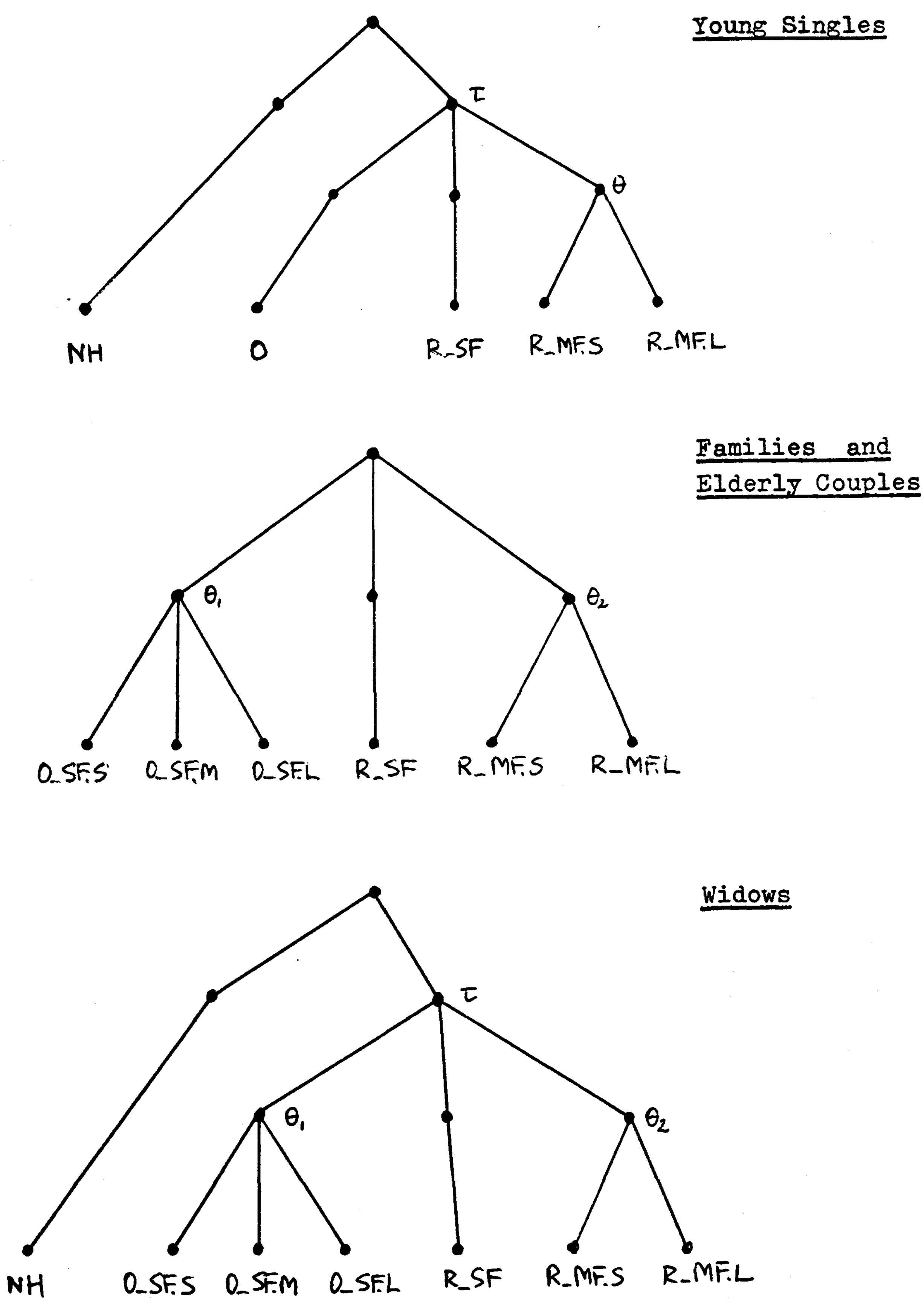


to dwellings with five or six rooms, and "large" to dwellings with at least seven rooms.

Some of the alternatives are fairly scarce, e.g., renting large apartments or single-family homes, so some alternatives have to be consolidated for a reliable estimation. This consolidation depends on the stratum. Furthermore, no cost data are available for owner-occupied dwellings in multifamily buildings, which forced us to omit these alternatives from the choice set. A more satisfactory approach would be to estimate the cost data for multifamily dwellings by hedonic regressions, or to explicitly model the missing alternatives in the definition of the choice probabilities. But the problem is a minor one for Dallas 1976 and Sacramento 1977, where these alternatives count for only 0.5 and 0.8 percent of all choices; it might bias our results only for Albany, where 6.9 percent of all nuclei chose cooperatively owned multi-family buildings. The final decision trees are depicted in Figure 2. (See Boersch-Supan and Pitkin (1982) for experiments with other decision trees.)

The choice among the alternatives will depend on the following demographic variables: (1) age of head of nucleus, (2) sex of head of nucleus, (3) marital status of head of nucleus, (4) race of head of nucleus, (5) number of children in the nucleus, as well as on financial variables of (6) user-cost and (7) income. We pursue the approach of de Leeuw (1971) and Quigley (1979) and assume different demand functions for nuclei with different demographic

Figure 2: Decision Trees for the Different Strata



characteristics. Accordingly, we stratify the sample with the first five demographic variables.

Of all possible strata, this paper examines demand functions for four representative strata:

(1) Young singles: unmarried white male and female without children, aged 20-35,
(2) Middle-aged families: married white couples with one or two children, aged 35-59,
(3) Elderly couples: married white couples without children, aged 60 and above,
(4) Widows: widowed, separated, and divorced white females without children, aged 60 and above.

The financial variables, income and user-cost, enter the demand functions directly. User-cost (UC) must be distinguished for renters and owners. For renters, the user-cost is simply gross rent. For owners, the user-cost has a number of components (see for example, Hendershott and Hu (1979) or Follain (1980)):

UC(own) = maintenance + insurance + utility-payments

+ mortgate-rate * debt

+ property tax rate * value

- tax savings from federal income tax deductions

+ T-Bill-rate * equity

- rate of appreciation * value

where the tax savings on the federal income tax is the sum of the local property tax and the mortgage interest, multiplied by the appropriate marginal tax rate. Note that federal income tax savings depend through the marginal tax rate on such nucleus characteristics as income and number of children. Furthermore, we assume different interest rates on debt and on equity to account for the effect of inflation on fixed-rate mortgages.

The user-cost of owners consists of two types of cost which the nucleus perceives differently: Maintenance, mortgage costs, property taxes and federal income tax savings are easily perceived as costs, whereas capital gains from appreciation are uncertain and opportunity costs of equity are a rather cloudy concept for non-economists. We therefore split up user-cost in two components:

UC(own) = OOPOCK(own) + RETURN(own),

where the "out-of-pocket cost" is composed of:

OOPOCK(own) = maintenance + insurance + utility-bills
+ local property tax + mortgage-rate * debt
- federal income tax savings,

and the return from the asset homeownership is defined as:

RETURN(own) = rate of appreciation * value
- T-Bill-rate * equity.

For fully rational housing demanders, the coefficients for OOPOCK and RETURN should be of equal magnitude and opposite sign. For renters, we set RETURN to zero. For households with more than one nucleus, some cost-sharing agreement has to be postulated: we assume that RETURN falls entirely to the head, whereas the out-of-pocket costs are shared according to the number of adults and number of children:

$$\text{SHARE(nucleus)} = \frac{\text{N\_ADULTS(nucleus)} + .5 * \text{N\_CHILDREN(nucleus)}}{\text{N\_ADULTS(household)} + .5 * \text{N\_CHILDREN(household)}}$$

The Annual Housing Survey gives us rather precise data for out-of-pocket costs. The return variable has to be constructed with external information: appreciation is based on the difference in house values between the Annual Housing Survey 1976/77 and the 1970 Census, converted into yearly rates. This rate varies by SMSA and by type of dwelling. Equity costs are calculated from the value-to-loan ratio in the Annual Housing Survey, multiplied by the interest on five-year U.S. treasury bills. Both appreciation rates and equity costs suffer from serious data problems: loan-to-value ratios are often not reported, making constructed substitutes necessary, and the available appreciation rates vary only by SMSA, but not within SMSA.

Given the cross-sectional data, the choice among the housing alternatives is a hypothetical one: we observe each nucleus with its chosen alternative and its attributes like user-cost, do not observe the attributes of the alternatives that the nucleus rejected. As a plausible approximation, we take as these attributes the averages in the cross-section confined to recent movers. For the hypothetical loan-to-value ratios, we assume a 20 percent downpayment for young singles, 50 percent for families, and 99 percent for the elderly households, which takes into account the availability of mortgage loans to the different age groups. This assumption is not critical to the estimates, but is confirmed by cross- tabulations of recent movers.

However, it is a rather critical assumption that prices of recent movers represent the hypothetical prices of not chosen units, because it essentially ignores the historical way the nucleus ended up in his housing choice with its current prices. This point reflects the main problem of this static model, which we will discuss in the final section of the paper.

Finally, income is defined as the current total gross income of all nucleus members. Due to the nature of a choice between discrete alternatives, income enters the demand functions interactively with alternative-specific dummies. Furthermore, income influences the out-of-pocket costs of homeowners because federal income tax savings depend on the marginal tax rate, i.e., on gross income.

Microeconomics and Econometrics of Hierarchical Choice Models

Let us assume the housing market is partitioned into M discrete housing alternatives, e.g., as depicted in Figure 1. We associate each of these alternatives with an index of desirability, which comprises all advantages and disadvantages for a given household (or, in this case, nucleus) into one scalar unit. In the language of neoclassical consumer theory, this is an index that orders all pairs (utility, price) of the M alternatives. Uncertainty about quality and erratic or irrational valuations introduce a stochastic component into this index.

Like the hypothesis of utility maximization under budget restriction, we assume that each household will choose the alternative with the highest index of desirability. Due to the probabilistic nature of the index, we will call this the random utility maximization hypothesis (McFadden, 1981). For each household t we decompose the desirability index u_{it} of the alternative i into a deterministic and a stochastic part:

$$u_{it} = v_{it} + e_{it}$$

The deterministic part is dependent on the characteristics of the alternative as well as on the characteristics of the household:

$$v_{it} = \sum_{k} x^{k}_{it} * b_{k} + \sum_{l} y^{l}_{t} * a_{l}$$

where x^k_{it} = the k-th characteristic of alternative i

for household t,

y^l_t = the l-th characteristic of household t,

a_l and b_k = weights (to be estimated).

In addition to uncertainty and erratic valuations, the stochastic disturbance e_{it} will pick up deviations of the household t from the weights a_l and b_k in the population. The different components of e_{it} can not be identified or only under specific assumptions.

Household t will choose alternative i, if

$$u_{it} > u_{jt} \quad \text{for all } j \neq i$$

Thus, the probability that household t chooses i among all M possible alternatives is

$$p_t(i) = \text{prob}(\ v_{it}+e_{it} > v_{jt}+e_{jt} \mid j \neq i\ )$$

$$= \int_{e_{it}=-\infty}^{\infty} \int_{e_{1t}=-\infty}^{u_{it}-v_{1t}} \cdots \int_{e_{Mt}=-\infty}^{u_{it}-v_{Mt}} dF(e_{1t},\ldots,e_{Mt})$$

where F denotes the joint cumulative distribution function of the errors e_{it}.

If the households t are a random sample of the population, the aggregation

$$f(i) = \sum_t p_t(i)$$

will yield the relative frequencies of alternative i in the population.

For a given specification of the deterministic utility v_{it}, the choice of a functional form for the relation between the choice probabilities $p_t(i)$ and the explanatory variables x^k_{it} and y^l_t is equivalent to the specification of the joint distribution F of the error terms e_{it}.

The integral formula shows the dilemma for this choice. On one hand, the correlation among the e_{it} should be as flexible as possible to allow different correlations among the choice probabilities. On the other hand, the computational effort of evaluating the multi dimensional integral should be minimized, suggesting a distribution function F where this can be done explicitly. This in particular prohibits the use of a normal distribution for problems with more than four alternatives.

Two families of distribution functions allow easy evaluation of the integral. One leads to a linear functional relation between the choice probabilities and the explanatory variables, and thus does not take account of the adding up and the unity interval restrictions of the choice probabilities. The other family is that of generalized extreme-value distributions, an extension of the logit approach; this is the family we will use to specify the choice probability.

A completely free correlation structure of the disturbances implies the estimation of $M*(M-1)/2$ correlation coefficients which is impractical for most sets of alternatives. Thus, further restrictions are necessary. The most drastic restriction is to postulate the independence of the e_{it}. Then the multi dimensional integral can be factorized, and the resulting choice probabilities are of the well-known multinomial logit form.

An application to the housing market can be found in Quigley (1976). The assumption of independent e_{it} is known as "Independence of Irrelevant Alternatives" (McFadden, 1973) due to the following necessary and sufficient characterizations:

(1) The odds of choosing alternative i over alternative j are independent of the attributes of all other alternatives and independent of the existence of any other alternative.
(2) The elasticity of the relative frequency f(i) of alternative i with respect to the attributes of any other alternative $j \neq i$ is constant, that is independent of j.

Therefore, independence can only be assumed for alternatives that are "equally different," but not for alternatives with different degrees of substitution. The following example translates a classical example (Domencich and McFadden, 1975) into the housing market. For simplicity, consider the tenure choice. Let us assume the relative odds are 1:1 for renting versus owning. Let us introduce a third, new form of tenure (e.g., cooperative) which is an almost ideal substitute for owning. Intuitively, we would expect the new distribution to be something like 50% : 25% : 25%. But condition (1) tells us that the relative odds of renting versus owning have to stay constant, forcing the new distribution to be 33% : 33% : 33%, which is implausible because of the similarity of owning and owning cooperatively.

The failure to accommodate different degrees of cross-alternative substitution renders the multinomial logit specification inappropriate for such heterogeneous choice sets as depicted in Figure 1. On the other hand, the possibility of grouping or clustering the alternatives according to their degree of substitution allows us a relatively straightforward way of combining the computational simplicity of the multinomial logit form with a richer substitution pattern: for each cluster, we introduce a parameter that describes the similarity of its alternatives. We can do the same with clusters themselves, and thereby achieve a hierarchical structure of similarities and substitution patterns. Within each cluster and between the clusters, we apply multinomial logit choice probabilities. This approach is called "Nested Multinomial Logit" (NMNL). (See McFadden, 1981 for a discussion of the development of these models.)

For the application at hand, let us introduce three steps of clustering. First, we bundle housing alternatives by size and quality, then these clusters by tenure and type of building, and finally by all headship alternatives versus the nonheadship alternative. Thus, we can look at NMNL models in another way and interpret them as hierarchical decision processes or decision trees as depicted in Figures 1 and 2, where each nucleus first decides whether to head a household or not, then the heads decide about tenure, and finally size. Note that this does not imply a temporal decomposition of the decision process, only a decomposition into classes of substitutability. See Boersch-Supan and Pitkin (1982) for a comparison of both views.

The choice probabilities for a three-level hierarchical decision process are composed of the conditional decisions at each level (we suppress the index t for the individual household):

$$p(i) = p_H(H_i) * p_T(T_i|H_i) * p_S(S_i|H_i,T_i)$$

where H_i = headship choice implied by alternative i,

T_i = tenure choice implied by alternative i,

S_i = size choice implied by alternative i.

At each level, the conditional choice probabilities have the multinomial logit form:

$$p_S(S_i|H_i,T_i) = \exp(v_i) \;/ \sum_j \exp(v_j),$$

with summation over all sizes j in tenure T_j

and v_i = index of desirability,

$$p_T(T_i|H_i) = \exp(\ c_i * I(T_i)\ ) / \sum_{T_j} \exp(\ c_j * I(T_j)\ ),$$

with summation over each tenure T_j in H_j

and $I(T_i) = \log \sum_j \exp(v_j)$ "inclusive value,"

$$p_H(H_i) = \exp(\ d_i * J(H_i)\ ) / \sum_{H_j} \exp(\ d_j * J(H_j)\ ),$$

with summation over both headship-possibilities H_j

and $J(H_i) = \log \sum_{T_j} \exp(c_j * I(T_j))$ "inclusive value."

Note the "similarity parameters" c_i and d_i which parametrize the degree of substitutability in each cluster. They can be interpreted as the regression coefficients for the "inclusive values" which are the equivalents of the desirability indices on the level of clusters and measure the surplus generated by all members in a cluster. If these parameters are one, the decision tree collapses to the multinomial logit model. If they are smaller than one, alternatives in the respective clusters are close substitutes relative to other alternatives, ideal substitutes in the case of zero. If all similarity parameters are in the unit-interval, the underlying joint distribution of the disturbances is well behaved and consistent with the microeconomic theory outlined at the beginning of this section, independent of the explanatory variables (see McFadden, 1978, 1979 and Daly and Zachary, 1979). With similarity parameters outside the unit-interval, this consistency will hold only for a certain range of explanatory variables, and it must be checked, whether this range includes the given data. This check and a potential reconciliation of such NMNL models with the random utility maximization hypothesis is discussed in the appendix.

Baseline Estimates

The sum of the logarithms of the choice probabilities in the preceding section is the likelihood function of the hierarchical choice model. Thus, the model can be estimated by maximizing over the taste weights a_l, b_k, and the similarity coefficients c_i, d_i. This can be done either sequentially by level of clustering as in Domencich and McFadden (1975) or Anas (1982), or jointly as in Coslett (1978) or Boersch-Supan and Pitkin (1982). Because the Full Information Maximum Likelihood function is highly nonlinear in the similarity parameters, the second approach is costly. However, the first approach is inefficient, especially for complex decision trees. We therefore prefer joint estimation and use the modified quadratic hill-climbing method developed by Goldfeld and Quandt (1972) with analytical first and numerical second derivatives. This procedure proved computationally fairly efficient compared with BHHH procedures (Berndt, Hall, Hall, Hausman, 1974).

The parameter estimates and summary statistics are tabulated in Tables 1-4 for each of the four strata. The parameters represent the taste weights of the respective explanatory variables in the deterministic part of the utility-function v_{it}. T-statistics are given in brackets, and are evaluated at zero for the taste weights. Note that income Y interacts with alternative specific dummies, where we use the same mnemonics for the alternatives as in Figure 1. The final parameters are the similarity parameters that express the degree of closeness in the respective clusters (See Figure 2). T-statistics

Table 1 : NMNL Parameter Estimates

"Young Singles": Unmarried, Age 20-35, No Children

	ALBANY	DALLAS	SACRAMENTO
OOPOCK	-0.69598	-1.18246	-1.56947
	(10.04)	(9.96)	(13.29)
RETURN	0.13156	0.12316	0.20421
	(1.55)	(1.45)	(2.97)
Y_NH	-0.02006	-0.13204	-0.14574
	(1.65)	(6.88)	(5.36)
Y_O	-0.00196	0.00894	0.02125
	(0.09)	(0.43)	(0.65)
Y_R_SF	0.01213	-0.02375	0.01408
	(1.87)	(1.61)	(0.62)
Y_R_MF.S	0.02282	0.04212	0.03539
	(4.07)	(3.38)	(1.59)
TH_R_MF	0.20422	0.42742	0.45242
	(26.81)	(8.05)	(7.08)
TAU_HEAD	0.14958	0.32322	0.37700
	(28.96)	(13.17)	(12.48)
LOGLIK	-616.591	-421.180	-457.465
LOGLIK_0	-1401.82	-925.427	-1199.03
RHO_SQ	0.56	0.545	0.618
%CORRECT	82.9%	73.2%	81.2%
NOBS	871	575	745

In brackets: t-statistics around zero or one
LOGLIK = loglikelihood at optimum
LOGLIK_0 = loglikelihood at zero
RHO_SQ = 1.0 - LOGLIK/LOGLIK_0
%CORRECT = percentage of correct ex post predictions
NOBS = number of observations

Table 2 : NMNL Parameter Estimates

"Families": Married, Age 35-59, 1-2 Children

	ALBANY	DALLAS	SACRAMENTO
OOPOCK	-0.92185	-3.55522	-2.81777
	(4.91)	(5.87)	(6.28)
RETURN	-0.69255	1.07524	0.63690
	(3.55)	(5.08)	(3.87)
Y_O_SF.S	-0.99077	-0.70474	-0.32266
	(4.35)	(4.69)	(1.91)
Y_O_SF.M	0.07448	0.05039	0.15820
	(3.12)	(1.20)	(1.21)
Y_O_SF.L	0.22973	0.20642	0.27165
	(6.19)	(3.56)	(2.08)
Y_R_SF	-0.00667	-0.02119	0.18802
	(0.37)	(0.63)	(1.49)
Y_R_MF.S	-0.10296	0.18602	-0.00061
	(1.86)	(1.69)	(0.00)
TH_R_MF	0.74920	2.25563	1.74344
	(0.71)	(0.86)	(1.06)
TH_O_SF	5.03934	2.73284	2.20262
	(8.77)	(3.21)	(2.97)
LOGLIK	-357.208	-121.134	-127.350
LOGLIK_0	-627.116	-584.114	-580.530
RHO_SQ	0.43	0.793	0.781
%CORRECT	78.9%	90.8%	90.9%
NOBS	350	326	324

In brackets: t-statistics around zero or one
LOGLIK = loglikelihood at optimum
LOGLIK_0 = loglikelihood at zero
RHO_SQ = 1.0 - LOGLIK/LOGLIK_0
%CORRECT = percentage of correct ex post predictions
NOBS = number of observations

Table 3 : NMNL Parameter Estimates

"Elderly Couples": Married, Age 60+, No Children

	ALBANY	DALLAS	SACRAMENTO
OOPOCK	-3.61354	-2.35330	-3.21796
	(8.29)	(6.66)	(6.49)
RETURN	0.52908	0.68032	1.02552
	(2.27)	(4.27)	(4.87)
Y_O_SF.S	-0.26210	-0.19193	-0.21797
	(4.15)	(2.05)	(2.68)
Y_O_SF.M	0.16994	0.08080	0.07719
	(3.96)	(1.02)	(1.11)
Y_O_SF.L	0.30284	0.25677	0.25455
	(4.94)	(2.67)	(3.05)
Y_R_SF	0.03169	0.03272	0.11807
	(0.59)	(0.42)	(1.61)
Y_R_MF.S	-0.08036	0.01262	-0.14165
	(1.17)	(0.15)	(1.01)
TH_R_MF	3.02968	2.08399	3.39293
	(2.93)	(1.49)	(1.85)
TH_O_SF	1.41709	0.87605	0.47555
	(1.36)	(0.49)	(1.73)
LOGLIK	-153.730	- 77.926	- 62.706
LOGLIK_0	-582.322	-458.690	-519.610
RHO_SQ	0.735	0.83	0.88
%CORRECT	86.8%	90.2%	91.7%
NOBS	325	256	290

In brackets: t-statistics around zero or one
LOGLIK = loglikelihood at optimum
LOGLIK_0 = loglikelihood at zero
RHO_SQ = 1.0 - LOGLIK/LOGLIK_0
%CORRECT = percentage of correct ex post predictions
NOBS = number of observations

Table 4 : NMNL Parameter Estimates

"Widows": Widowed, Divorced, Separated, Age 60+, No Children

	ALBANY	DALLAS	SACRAMENTO
OOPOCK	-2.92313	-1.74357	-3.65622
	(9.81)	(7.17)	(7.61)
RETURN	0.28555	0.16035	0.61589
	(2.65)	(2.17)	(4.84)
Y_NH	-0.68624	-0.49553	-0.83475
	(10.42)	(4.77)	(6.47)
Y_O_SF.S	-0.23088	-0.16355	-0.27639
	(4.71)	(1.95)	(3.12)
Y_O_SF.M	0.22788	0.09891	0.16234
	(6.47)	(1.36)	(1.99)
Y_O_SF.L	0.39921	0.20989	0.43735
	(5.73)	(2.09)	(2.88)
Y_R_SF	0.01295	0.06526	0.01310
	(0.34)	(0.89)	(0.21)
Y_R_MF.S	-0.02053	0.04518	-0.00102
	(0.63)	(0.59)	(0.01)
TH_O_SF	1.48636	0.60497	1.03269
	(1.75)	(2.10)	(0.07)
TH_R_MF	1.48222	1.48958	0.97076
	(1.73)	(1.12)	(0.09)
TAU_HEAD	0.48036	0.45743	0.47043
	(5.79)	(6.31)	(4.21)
LOGLIK	-311.114	-199.974	- 94.920
LOGLIK_0	-1037.17	-646.042	-651.880
RHO_SQ	0.70	0.69	0.85
%CORRECT	89.0%	88.3%	82.6%
NOBS	533	332	335

In brackets: t-statistics around zero or one
LOGLIK = loglikelihood at optimum
LOGLIK_0 = loglikelihood at zero
RHO_SQ = 1.0 - LOGLIK/LOGLIK_0
%CORRECT = percentage of correct ex post predictions
NOBS = number of observations

for the similarity parameters are evaluated at one, using the multinomial logit case as a benchmark. Three scalar measures of performance or fit are used. The straightforward discrete analogy to the continuous R^2, using the sum of squared errors, has no discriminatory power for the model at hand: it is .99 for almost all strata. A more satisfactory measure is the ratio of the likelihood at the estimated parameters and the likelihood with taste weights at zero and similarity parameters at one. One minus this ratio behaves like the continuous R^2 (see McFaddden, 1973 or Amemiya, 1981; Domencich and McFadden (1975) give a comparison between the latter two measures of fit and their discriminatory power). As a third measure of fit, we compare actual with predicted individual choices which is a fairly stringent, though erratic criterion. Note that discrete choice models produce two predictions of the aggregate choice probabilities:

$$(1)\quad f(i) = \sum_t p_t(i)$$

$$(2)\quad f(i) = n(i)/T$$

$$\text{where } n(i) = \text{number } \{ p(i) = \max p(j) \mid j=1...M \}$$

$$= \text{number of people who chose alternative i}$$

$$T = \text{sample size}$$

The erratic nature of the percentage of correct predictions is due the integer constraint in (2). Table 5 gives an example of a success table in which observed and predicted alternatives are compared. The off-diagonal elements show the mispredictions: in this case, the model has some difficulties in discriminating between small rental

Table 5: Example of a Success Table and a Full Elasticity Matrix

PREDICTION SUCCESS TABLE:

OBSERVED ALT.	PREDICTED ALTERNATIVE NH	O_SF.S	O_SF.M	O_SF.L	R_SF	R_MF.S	R_MF.L
NH	73	1	0	0	0	10	0
O_SF.S	1	32	0	0	0	1	0
O_SF.M	0	0	77	0	0	1	0
O_SF.L	0	0	0	19	0	0	0
R_SF	2	0	0	0	25	1	0
R_MF.S	17	0	0	0	1	70	0
R_MF.L	1	0	0	0	0	1	2

PERCENT CORRECTLY PREDICTED : 88.96 %

MEAN INDIVIDUAL ELASTICITIES:

VARIABLE	ALT.	CHOICE PROBABILITY OF ALTERNATIVE: NH	O_SF.S	O_SF.M	O_SF.L	R_SF	R_MF.S	R_MF.L
OOPOCK	NH	-2.814	0.739	0.739	0.739	0.739	0.739	0.739
OOPOCK	O_SF.S	0.420	-12.898	-4.671	-4.671	1.876	1.876	1.876
OOPOCK	O_SF.M	0.918	0.389	-12.540	0.389	1.968	1.968	1.968
OOPOCK	O_SF.L	0.235	0.195	0.195	-23.002	0.502	0.502	0.502
OOPOCK	R_SF	0.287	0.670	0.670	0.670	-19.015	0.670	0.670
OOPOCK	R_MF.S	1.315	4.097	4.097	4.097	4.097	-10.477	-3.161
OOPOCK	R_MF.L	0.103	0.322	0.322	0.322	0.322	-0.317	-10.964
RETURN	O_SF.S	-0.200	2.759	0.892	0.892	-0.634	-0.634	-0.634
RETURN	O_SF.M	-0.429	-0.345	1.711	-0.345	-0.915	-0.915	-0.915
RETURN	O_SF.L	-0.120	-0.112	-0.112	2.316	-0.256	-0.256	-0.256
Y_NH		-4.422	0.588	0.588	0.588	0.588	0.588	0.588
Y_O_SF.S		0.187	-2.059	-0.453	-0.453	0.452	0.452	0.452
Y_O_SF.M		-0.330	-0.242	0.702	-0.242	-0.704	-0.704	-0.704
Y_O_SF.L		-0.302	-0.281	-0.281	2.261	-0.642	-0.642	-0.642
Y_R_SF		-0.005	-0.012	-0.012	-0.012	0.155	-0.012	-0.012
Y_R_MF.S		0.001	0.004	0.004	0.004	0.004	-0.009	-0.003
Y_SUM		-4.871	-2.002	0.548	2.146	-0.147	-0.326	-0.320

housing and non-headship, which might be due to the relatively crude specification of the household formation process.

The model achieves a surprisingly high prediction accuracy in terms of all three measures of fit. This is surprising because of the small number of explanatory variables and the simple specification. The model performs poorest in the strata of young singles and in the family stratum in Albany. The first is not astonishing: a static model can hardly capture changes in housing consumption in this period when the nucleus is establishing its own existence. These strata are also very heterogenous and include children still living with their parents, student roommates, and singles in their thirties. The poor performance in the case of families with one or two children in Albany may be attributable to the misspecification of the decision tree where the alternative of owning cooperatively is not included.

The main result is the significance of the price varibles in all strata. The out-of-pocket costs are highly significant, while the RETURN variable is somewhat weaker. Note that the hypothesis of rationality -- i.e., equal magnitude and opposite signs for the taste weights of OOPOCK and RETURN -- is rejected; considerably more weight is given to easily perceived out-of-pocket costs as opposed to appreciation minus equity costs. One should keep in mind, however, the difficulties of constructing the RETURN variable. Note furthermore that RETURN is least significant for the young singles, the strata most affected by liquidity constraints, rendering the rationality hypothesis inappropriate and introducing a lot of noise.

Note the similarity between the estimates for elderly couples and elderly widows: taking into account the different choice sets (a large proportion of widows live in their children's homes), it reflects the stability of the taste weights during old age.

The income dummies have a threefold function. First, they reflect the relative price of housing with respect to all other goods. In addition, they indicate the attractiveness of the various alternatives relative to large rented apartments, measured in money terms. In absence of any other alternative-specific dummies, they also pick up all other unmeasured advantages and disadvantages of the included alternatives relative to large rented apartments. Thus, one should be careful not to rush to conclusions about pure income effects. Introduction of alternative-specific dummies in several test strata reduces the income parameters, but leaves the price variables virtually constant. Because the focus of the simulations is on relative prices rather than on income, we avoided the costly inclusion of alternative specific dummies.

The attractiveness of the alternatives measured by the taste weights of the income dummies corresponds to *a priori* assessment. Note that most of the rented single-family houses are small houses, thus their negative weight for families with one or two children.

The last two coefficients in Tables 1-4 are the weights of the inclusive values or similarity coefficients. Note that four of the similarity coefficients are significantly larger than one (at the 5 percent level). This implies that in these strata the compatibility with the underlying microeconomic theory of random utility maximization must be explicitly checked for the given data and is not automatically guaranteed as in the other strata (see appendix). In fact, the test rejects this compatibility. Note that the microeconomic theory described above is based on static utility maximization. Note furthermore that failure of the test occurs in the strata where people move considerably less frequently than in the strata of young singles, where the similarity parameters are in the unit-interval. The rejection thus could be interpreted as a hint that optimization is done dynamically and that the model in these strata should be interpreted as only a reduced-form description of the steady-state as opposed to a structural static choice model.

The taste weights in Tables 1-4 can be transformed into elasticities of the choice probabilities with respect to the various explanatory variables:

$$\frac{d\,p(i)}{d\,x_{jk}} = a_k * x_{jk} * \Big(-p(j) + k_S * 1/c + k_T * (1/c-1/d) * Q(S) + k_H * (d-1)/d * Q(S) * Q(T) \Big)$$

where $k_S = 0$ if i and j are in the same size category
$= 1$ otherwise

$k_T = 0$ if i and j are in the same tenure category
$= 1$ otherwise

$k_H = 0$ if i and j are in the same headship category
$= 1$ otherwise

c,d = similarity parameters: $c=c_T$ and $d=d_H$

$Q(S)$ = conditional choice probability $p_S(S_j|H_j,T_j)$

$Q(T)$ = conditional choice probability $p_T(T_j|H_j)$

Note that for the cross elasticities the difference between i and j enters only through the "switches" k_S, k_T, and k_H. The structure of the tree is therefore directly reflected in the pattern of cross elasticities. This can be seen by comparing Figure 2 with the full matrix of cross elasticities in Table 5.

Derived from a highly non-linear model, elasticities at variable means are generally different from mean individual elasticities. Own price elasticities and income elasticities, tabulated in Table 6, refer to a change of the probability of choosing alternative i, when OOPOCK or RETURN in alternative i is changed. The income elasticities in Table 6 are the sum over the elasticities of all income dummies. (See Table 5 for an example of the individual elasticities.)

In interpreting the elasticities, one should keep the choice-probabilities and their nonlinearity in mind (See Table 7); the elasticities tend to be very high at very low probabilities and vice versa, reflecting saturation effects. Thus, comparisons among strata should be made with care. As a general pattern, the strata of young singles and elderly widows are the most price-responsive, especially in the owner alternatives, reflecting a priori knowledge of inertia and mobility in the different strata. Return from the asset homeownership exhibits a strong life-cycle behavior, and is thus higher for young people with a long decision horizon than for the elderly. Headship rates are highly responsive to prices for both

Table 6: Own Price and Sum of Income Elasticities

PROB	NH	O_SF.S	O_SF.M	O_SF.L	R_SF	R_MF.S	R_MF.L
Albany, Young Singles:							
OOPOCK	-0.395		-19.596		-12.973	-2.463	-7.227
RETURN	0.0		3.675		0.0	0.0	0.0
INCOME	-0.083		-0.591		-0.101	0.238	-0.343
Albany, Families							
OOPOCK		-0.209	-0.686	-0.898	-2.460	-2.517	-2.169
RETURN		-0.468	-0.386	-0.415	0.0	0.0	0.0
INCOME		-4.835	-0.132	0.553	-3.394	-6.250	-3.193
Albany, Elderly Couples:							
OOPOCK		-3.765	-6.222	-11.998	-10.187	-4.517	-5.160
RETURN		0.791	0.379	0.653	0.0	0.0	0.0
INCOME		-4.214	0.028	1.333	-1.788	-2.944	-2.574
Albany, Widows:							
OOPOCK	-2.430	-6.984	-7.529	-11.861	-16.940	-5.979	-6.016
RETURN	0.0	-0.914	-0.292	-0.418	0.0	0.0	0.0
INCOME	-3.694	-1.284	0.359	0.973	-0.428	-0.737	-0.664
Dallas, Young Singles:							
OOPOCK	-0.632		-16.494		-8.392	-3.031	-7.641
RETURN	0.0		2.165		0.0	0.0	0.0
INCOME	-0.669		-0.037		-0.710	0.625	-0.030
Dallas, Families:							
OOPOCK		-2.167	-3.366	-6.141	-8.548	-4.786	-8.590
RETURN		1.531	1.335	1.451	0.0	0.0	0.0
INCOME		-7.663	-1.343	1.138	-2.536	-4.690	-2.513
Dallas, Elderly Couples:							
OOPOCK		-4.581	-5.189	-18.934	-5.798	-4.316	-4.311
RETURN		2.064	1.313	2.497	0.0	0.0	0.0
INCOME		-5.462	-0.667	2.427	-1.337	-1.676	-1.770
Dallas, Widows:							
OOPOCK	-1.431	-5.806	-8.680	-24.426	-8.964	-5.162	-4.867
RETURN	0.0	0.816	0.682	1.016	0.0	0.0	0.0
INCOME	-2.823	-1.958	0.457	1.478	0.341	0.005	-0.164

Table 5: Own Price and Sum of Income Elasticities (cont'd)

PROB	NH	O_SF.S	O_SF.M	O_SF.L	R_SF	R_MF.S	R_MF.L
Sacramento, Young Singles:							
OOPOCK	-0.728		-16.459		-10.046	-3.563	-9.519
RETURN	0.0		2.939		0.0	0.0	0.0
INCOME	-0.513		0.279		0.181	0.468	0.064
Sacramento, Families:							
OOPOCK		-2.570	-2.862	-4.790	-6.922	-5.222	-5.358
RETURN		1.215	0.844	1.112	0.0	0.0	0.0
INCOME		-5.563	-0.550	0.632	-0.397	-4.727	-4.719
Sacramento, Elderly Couples:							
OOPOCK		-13.100	-9.361	-42.173	-8.112	-4.318	-5.485
RETURN		5.479	2.490	8.194	0.0	0.0	0.0
INCOME		-7.813	0.672	5.770	0.765	-2.055	-1.484
Sacramento, Widows:							
OOPOCK	-2.814	-12.898	-12.540	-23.002	-19.015	-10.477	-10.964
RETURN	0.0	2.759	1.711	2.316	0.0	0.0	0.0
INCOME	-4.871	-2.002	0.548	2.146	-0.147	-0.326	-0.320

young singles and elderly widows. Finally, note again that the income elasticities measure not only income but also pure alternative specific effects due to their interaction with alternative specific dummies.

The elasticity pattern is fairly stable across the three SMSAs, in spite of their very different distribution of housing alternatives. This provides some confidence in the robustness of the model. As a general pattern, housing demand reacts most to prices in Sacramento and least in Albany, suggesting the more flexible nature of the housing market in California compared with New England. The pattern holds for both out-of-pocket costs and returns.

Summing up, we observe the following:

o Relative prices significantly determine housing choices for given demographic variables.
o Household formation, in particular, is highly responsive.
o Out-of-pocket costs have higher taste weights than return from homeownership.
o Among strata, young singles and widows are more price responsive than the relatively inert strata of families and elderly couples.
o The sensitivity to RETURN shows the expected life-cycle behavior.
o The general pattern of elasticities is fairly stable across markets, with Albany behaving least flexibly and Sacramento the most.

A Housing Allowance Program

Between 1973 and 1979, the U.S. Department of Housing and Urban Development conducted a large scale Experimental Housing Allowance Program. Kennedy (1980) describes in detail the design of the program, and a good survey of the subsequent discussion and critique is given in Bradbury and Downs (1981). Somewhat surprising is the fact that all components of the Experimental Housing Allowance Program ignored the feedback of housing allowances on household formation. One focal point in this section is the question of how much improvement in housing conditions comes through increased headship rates over and above moves of existing households into larger dwellings.

The following simulation assumes a so-called housing gap formula for the calculation of the allowances. First, for each family size and site a benchmark rent is calculated, representing the "fair cost of standard housing." Then a minimum standard of quality is established, with only dwellings above this standard eligible for the subsidy. Finally, a linear tax is levied on the allowances in such a way that people with no (adjusted) income will receive the full rent for standard housing, whereas people above a certain income level will receive no allowances at all.

If the minimum standard is measured as a fraction a of the fair cost of standard housing C, and the upper income limit is a multiple b of C, then the housing allowance for a household with income Y and rent R is:

$$
\begin{array}{ll}
0 & \text{if } R < aC \\
0 & \text{if } Y > bC \\
C-Y/b & \text{otherwise.}
\end{array}
$$

To perform a realistic experiment, we use the settings a=0.7, b=4.0, and C from the Experimental Housing Allowances Program, where C was taken from the Pittsburgh demand experiment and inflated by a yearly as well as inter-SMSA rent index as follows:

FAIR MONTHLY RENTS NO. OF PERSONS:	PITTSBG 75	DALLAS 77	ALBANY 77	SACRAM 76
1	$ 115	$ 150	$ 130	$ 140
2	130	180	160	170
3-4	150	200	180	190
5-6	170	225	205	215
7+	205	275	245	260

Housing allowances introduce nonlinearities in the budget set, see Hausman and Wise (1980) or Venti and Wise (1982). They can be handled fairly elegantly in discrete choice models by changing the prices of the housing alternatives differently rather than by adding the allowances to the income.

Table 7 lists the predicted shares of the housing alternatives before and after the introduction of the housing allowance program. Given the static nature of the model, this reflects a change between steady-states. The shares are calculated as means of the individual

Table 7: HOUSING ALLOWANCES

Stratum	Alt.	ALBANY		DALLAS		SACRAMENTO	
YSL	NH	.6331	.4780	.5507	.3226	.6262	.3067
	O_SM						
	O_ME	.0092	.0092	.0224	.0219	.0425	.0409
	O_LA						
	R_SF	.0123	.0159	.0478	.0913	.0508	.0772
	R_SM	.2815	.4041	.3438	.5046	.2642	.5396
	R_LA	.0637	.0928	.0353	.0596	.0163	.0356
FAM	NH						
	O_SM	.0182	.0149	.0440	.0272	.0320	.0210
	O_ME	.3310	.3270	.4439	.4370	.4772	.4682
	O_LA	.4997	.4969	.4085	.4046	.3852	.3831
	R_SF	.0546	.0572	.0639	.0694	.0780	.0772
	R_SM	.0161	.0207	.0152	.0299	.0219	.0394
	R_LA	.0804	.0833	.0244	.0320	.0056	.0111
ELC	NH						
	O_SM	.1251	.0920	.1443	.1174	.1450	.1191
	O_ME	.4050	.3830	.5446	.5316	.6667	.6418
	O_LA	.2773	.2683	.2014	.2004	.0750	.0746
	R_SF	.0174	.0193	.0420	.0517	.0531	.0595
	R_SM	.0870	.1225	.0484	.0686	.0335	.0613
	R_LA	.0882	.1149	.0192	.0302	.0268	.0437
WID	NH	.2757	.1300	.2130	.1005	.2745	.1068
	O_SM	.0694	.0387	.1165	.0562	.1105	.0946
	O_ME	.1602	.1398	.3144	.2875	.2281	.2232
	O_LA	.0985	.0909	.0486	.0453	.0562	.0558
	R_SF	.0273	.0286	.0737	.0953	.0729	.0777
	R_SM	.2340	.3636	.1842	.3039	.2437	.4160
	R_LA	.1348	.2084	.0496	.1133	.0140	.0259

First column : predicted shares of housing alternatives before housing allowances.

Second column: predicted shares of housing alternatives with housing allowance program in effect (housing gap formula: $P = C - Y/b$).

choice probabilities. Table 8, in turn, tabulates the moves according to the individual predictions.

Our main result is the strong impact of housing allowances on headship rates: about half of the people who lived in some sort of shared accommodations created their own households in response to the housing allowance program. Most of these nuclei in the strata of young singles have little or no income, thus their rent net of the housing allowance is virtually zero. More surprising is the strong response in the strata of elderly widows, where the non-head share is far less and the income higher than among the young singles, the share of non-heads is nevertheless drastically reduced in response to the subsidy. We conclude once more that headship rates are important endogenous variables in the housing market.

Within the rental sector, only few moves occur. The mobility rates induced by the housing allowances (Albany 0.047, Dallas 0.057, Sacramento 0.055) are very close to those measured in the demand part of the Experimental Housing Allowance Program by MacMillan (1980) i.e., Pittsburgh 0.045 and Phoenix 0.101. Note again the difference in the price sensitivity between the Northeast and the Southwest.

Unlike the Experimental Housing Allowance Program, our simulation offered allowances for rental housing to everybody in the population, changing the balance in the tenure choice in favor of renting. As a response, we observe a relatively large number of moves from the owner-occupied section into the rental section of the housing market.

Table 8: Actual Moves in Response to a Housing Allowance Program

Stratum	from	NH	O.S	O.M	O.L	R.SF	R.MS	R.ML
Albany,	to R_SF	1		0		-	0	0
Young Singles	to R_MF.S	309		0		0	-	0
	to R_MF.L	3		0		0	0	-
Albany,	to R_SF		0	0	0	-	0	0
Families	to R_MF.S		9	3	0	3	-	3
	to R_MF.L		0	0	3	0	3	-
Albany,	to R_SF		0	0	0	-	0	0
Elderly Couples	to R_MF.S		24	28	3	0	-	0
	to R_MF.L		6	0	0	0	0	-
Albany,	to R_SF	0	0	0	0	-	0	0
Widows	to R_MF.S	156	11	26	11	0	-	4
	to R_MF.L	0	0	0	2	0	2	-
Dallas,	to R_SF	4		0		-	0	0
Young Singles	to R_MF.S	426		0		0	-	0
	to R_MF.L	2		0		0	0	-
Dallas,	to R_SF		3	0	0	-	0	0
Families	to R_MF.S		18	6	0	3	-	0
	to R_MF.L		0	0	0	0	0	-
Dallas,	to R_SF		8	0	0	-	0	0
Elderly Couples	to R_MF.S		35	0	0	4	-	0
	to R_MF.L		0	0	0	0	0	-
Dallas,	to R_SF	18	0	0	0	-	0	0
Widows	to R_MF.S	172	33	12	3	12	-	0
	to R_MF.L	0	0	0	0	0	0	-
Sacramento,	to R_SF	9		0		-	0	0
Young Singles	to R_MF.S	525		1		3	-	1
	to R_MF.L	1		0		0	0	-
Sacramento,	to R_SF		3	0	0	-	0	0
Families	to R_MF.S		16	6	3	9	-	0
	to R_MF.L		0	0	0	0	0	-
Sacramento,	to R_SF		7	0	0	-	0	0
Elderly Couples	to R_MF.S		31	17	0	3	-	0
	to R_MF.L		0	0	0	0	0	-
Sacramento,	to R_SF	0	0	0	0	-	0	0
Widows	to R_MF.S	194	3	3	0	3	-	0
	to R_MF.L	0	0	0	0	0	0	-

Notes: predicted moves, normalized for 1000 nuclei per stratum

Table 9: INCIDENCE: BETWEEN JURISDICTIONS AND STRATA

Stratum	Level of Government	Housing Allowances	Property Tax Cut	Fed. Income Tax Change
Albany, Young Singles:				
	Federal, direct subsidy	638.60	0.0	0.0
	Federal, tax-subsidy	0.0	-1.02	-0.91
	Local, lost property tax	0.0	5.20	0.0
Albany, Families				
	Federal, direct subsidy	50.16	0.0	0.0
	Federal, tax-subsidy	0.0	-110.4	-33.73
	Local, lost property tax	12.30	433.9	1.94
Albany, Elderly Couples				
	Federal, direct subsidy	60.52	0.0	0.0
	Federal, tax-subsidy	-0.98	-77.54	-26.83
	Local, lost property tax	44.31	357.57	9.05
Albany, Widows				
	Federal, direct subsidy	274.65	0.0	0.0
	Federal, tax-subsidy	-0.33	-14.24	-6.28
	Local, lost property tax	39.37	80.87	3.81
Dallas, Young Singles				
	Federal, direct subsidy	823.08	0.0	0.0
	Federal, tax-subsidy	0.0	-0.79	-0.63
	Local, lost property tax	0.0	3.73	0.0
Dallas, Families				
	Federal, direct subsidy	46.24	0.0	0.0
	Federal, tax-subsidy	-0.91	-87.48	-43.63
	Local, lost property tax	10.93	286.66	3.73
Dallas, Elderly Couples				
	Federal, direct subsidy	58.22	0.0	0.0
	Federal, tax-subsidy	0.0	-47.80	-16.80
	Local, lost property tax	12.12	220.49	0.0
Dallas, Widows				
	Federal, direct subsidy	480.17	0.0	0.0
	Federal, tax-subsidy	0.0	-10.31	-4.21
	Local, lost property tax	19.58	123.76	0.0

Table 9: INCIDENCE: BETWEEN JURISDICTIONS AND STRATA (cont'd)

Stratum	Level of Government	Housing Allowances	Property Tax Cut	Fed. Income Tax Change
Sacramento, Young Singles				
	Federal, direct subsidy	982.60	0.0	0.0
	Federal, tax-subsidy	-0.04	-3.71	-1.60
	Local, lost property tax	1.92	9.72	0.67
Sacramento, Families				
	Federal, direct subsidy	45.74	0.0	0.0
	Federal, tax-subsidy	-2.77	-80.80	-38.42
	Local, lost property tax	13.70	338.09	7.31
Sacramento, Elderly Couples				
	Federal, direct subsidy	70.02	0.0	0.0
	Federal, tax-subsidy	-1.33	-56.72	-41.30
	Local, lost property tax	27.71	289.92	6.66
Sacramento, Widows				
	Federal, direct subsidy	299.39	0.0	0.0
	Federal, tax-subsidy	0.0	-13.78	-0.71
	Local, lost property tax	2.96	101.39	0.0

Notes: The table lists the direct subsidy in the case of housing allowances, the indirect subsidy via Federal Income Tax savings due to deduction of interest and local property tax, and the local property tax losses to the local jurisdiction. The unit is $ 1000 for a normalized stratum of 1000 nuclei, i.e. dollars per nucleus per year. The numbers are based on the predicted moves of Table 8.

The mobility rates for the shift from owning to renting induced by the allowances are between 0.125 for Sacramento and 0.189 for Albany. Note the lower rate for Sacramento, reflecting the high valuation of owner-occupancy in the West relative to the Northeast.

Moves from owner-occupancy into the rental market have two important fiscal side-effects: on the federal level, some money given for housing allowances is retrieved through lower mortgage and property tax deductions from the federal income tax. More important is the spill-over effect at the local level of reductions in local property taxes. Table 9 lists these fiscal repercussions. All amounts are normalized to a stratum of 1000 nuclei to allow for comparisons both among strata and among SMSA. Note that especially for the married strata, the losses in local property taxes are a sizable proportion of the housing allowances paid by the federal government.

We can sum up the results of the housing allowance experiment as follows:

- o headship rates are highly responsive to the housing subsidies,
- o mobility rates within the rental market are low and of comparable size to the findings of the Experimental Housing Allowance Program,
- o greater mobility between renting and owning produces sizable spill-over effects from federal policy to the local level.

Cutting the Local Property Tax By One Half

In recent years, some states have passed legislation that introduces upper ceilings for local property tax rates, e.g., Proposition 13 in California and Propostion 2-1/2 in Massachusetts. These ceilings imply a drastic reduction in local property taxes for given assessment ratios. As a crude approximation of the isolated impact due to a drastic change in the local property tax rate, the following simulation predicts the distribution of nuclei into housing categories assuming a property tax rate of half the level in effect during the estimation period 1976/77.

Effective property taxes (as percentages of the house values reported in the Annual Housing Survey) in this period were 2.2 percent in Albany, 1.3 percent in Dallas, and 1.6 percent in Sacramento. The proportion of property taxes in the out-of-pocket cost varies considerably across strata, mainly due to the variation in mortgage payments in the life cycle, and less so across housing alternatives; the overall proportion is about 10 percent. The impact of the property tax cut is softened by a reduction in the federal income tax deductions proportional to the marginal tax rate of the household. Taking this into account, the simulation reduces the cost of owner-occupancy about 3 percent for the average homeowner. This is a fairly small change in relative prices considering that the property tax rate is lowered by 50 percent.

Table 10 lists the distribution of housing alternatives before

Table 10 : LOCAL PROPERTY TAX

Stratum	Alt.	ALBANY		DALLAS		SACRAMENTO	
YSL	NH	.6331	.6322	.5507	.5501	.6262	.6227
	O_SM						
	O_ME	.0092	.0142	.0224	.0248	.0425	.0544
	O_LA						
	R_SF	.0123	.0119	.0478	.0477	.0508	.0494
	R_SM	.2815	.2791	.3438	.3424	.2642	.2578
	R_LA	.0637	.0627	.0353	.0350	.0163	.0157
FAM	NH						
	O_SM	.0182	.0201	.0440	.0458	.0320	.0378
	O_ME	.3310	.3413	.4439	.4585	.4772	.4899
	O_LA	.4997	.5123	.4085	.4134	.3852	.3940
	R_SF	.0546	.0456	.0639	.0530	.0780	.0587
	R_SM	.0161	.0132	.0152	.0114	.0219	.0157
	R_LA	.0804	.0674	.0244	.0178	.0056	.0038
ELC	NH						
	O_SM	.1251	.1745	.1443	.1529	.1450	.1617
	O_ME	.4050	.4173	.5446	.5503	.6667	.6728
	O_LA	.2773	.2826	.2014	.2023	.0750	.0770
	R_SF	.0174	.0137	.0420	.0374	.0531	.0436
	R_SM	.0870	.0543	.0484	.0412	.0335	.0246
	R_LA	.0882	.0576	.0192	.0159	.0268	.0203
WID	NH	.2757	.2396	.2130	.2016	.2745	.2585
	O_SM	.0694	.1606	.1165	.1411	.1105	.1506
	O_ME	.1602	.1746	.3144	.3202	.2281	.2326
	O_LA	.0985	.1011	.0486	.0494	.0562	.0572
	R_SF	.0273	.0254	.0737	.0703	.0729	.0711
	R_SM	.2340	.1942	.1842	.1703	.2437	.2181
	R_LA	.1348	.1046	.0496	.0450	.0140	.0120

First column : predicted shares of housing alternatives under actual 1977 local property taxes.

Second column: predicted shares of housing alternatives under only 50% of the 1977 local property taxes.

and after the property tax change, calculated as means of the individual choice probabilities. If we concentrate only on the tenure choice, the share of owner-occupancy increases by:

Stratum	Albany	Dallas	Sacramento
Young Singles	.0050	.0024	.0119
Families	.0249	.0213	.0273
Eldery Couples	.0670	.0151	.0249
Widows	.1080	.0333	.0454

The impact is of course strongest in Albany where the property tax is substantially higher than in Dallas and Sacramento. In addition, the impact is very low for young singles: they have high mortgage payments and the percentage of property taxes in their total out-of-pocket costs is very low. The same reasoning explains why the increase in owner-occupancy is largest for small houses. In addition, smaller houses are attractive for people with low incomes, for whom the offsetting effect of decreasing income tax deductions is least.

Finally, we can see the interjurisdictional fiscal effects in the second column of Table 9. In the family strata, the gains for the federal government by smaller deductions are between 23.9 percent and 30.3 percent of the losses in local property taxes. The size of this spill-over effect depends on two factors: it simply reflects the relatively high marginal tax rates for these strata, but the gains are also reduced by the higher share of owner-occupancy in response to the tax change.

We can summarize the results of the property tax experiment as follows:

- o The impact of a strong reduction in the local property tax is small in the strata with high mortgage payments and high tax brackets. It is high for the elderly and for small homeowners.
- o The spill-over effect to the federal government is sizable. The direct effect through the marginal tax rate is partially offset by the indirect effect of movers into owner-occupancy.

Making the Federal Income Tax Less Progressive

The final simulation concerns the change in the federal income tax law that reduced the highest marginal tax rate from 70 percent to 50 percent. This has two opposing effects on housing consumption: while high-income people pay fewer taxes the deductions for mortgage interest and local property taxes are less worth and thus reduce the tax advantages of ownership. In the following simulation, we isolate the second effect by holding the income level constant and calculate the tax savings in the out-of-pocket costs of homeownership assuming the new tax schedule. We used the federal income tax schedule for 1983 and deflated the tax brackets by the Consumer Price Index to the price and income level of the estimation period.

We can again make the back-on-the-envelope calculation as in the preceding section: for the very rich, deductions lose 20 percent of their value. If we assume that a third of the out-of-pocket costs is deductible, we generate a 7 percent increase in the cost of owner-occupancy. This is an upper limit: people in lower tax brackets face a much smaller increase because below the 50 percent brackets, the marginal tax rates were only very slightly reduced. For the poor, there is no change whatsoever.

It should be noted that the sample includes only few "very rich" people (the 50 percent tax bracket in 1977 was about $ 40,000) because the selection of strata overrepresents the very young and elderly nuclei. Table 11 shows that the change in the marginal tax rate

Table 11 : FEDERAL INCOME TAX

Stratum	Alt.	ALBANY		DALLAS		SACRAMENTO	
YSL	NH	.6331	.6333	.5507	.5508	.6262	.6269
	O_SM						
	O_ME	.0092	.0088	.0224	.0219	.0425	.0408
	O_LA						
	R_SF	.0123	.0124	.0478	.0478	.0508	.0509
	R_SM	.2815	.2817	.3438	.3441	.2642	.2651
	R_LA	.0637	.0638	.0353	.0353	.0163	.0163
FAM	NH						
	O_SM	.0182	.0179	.0440	.0427	.0320	.0297
	O_ME	.3310	.3284	.4439	.4400	.4772	.4734
	O_LA	.4997	.4970	.4085	.4077	.3852	.3833
	R_SF	.0546	.0566	.0639	.0666	.0780	.0830
	R_SM	.0161	.0166	.0152	.0168	.0219	.0242
	R_LA	.0804	.0835	.0244	.0262	.0056	.0064
ELC	NH						
	O_SM	.1251	.1210	.1443	.1440	.1450	.1438
	O_ME	.4050	.4034	.5446	.5433	.6667	.6645
	O_LA	.2773	.2760	.2014	.2012	.0750	.0744
	R_SF	.0174	.0177	.0420	.0425	.0531	.0542
	R_SM	.0870	.0905	.0484	.0493	.0335	.0352
	R_LA	.0882	.0914	.0192	.0193	.0268	.0280
WID	NH	.2757	.2760	.2130	.2131	.2745	.2747
	O_SM	.0694	.0676	.1165	.1163	.1105	.1108
	O_ME	.1602	.1596	.3144	.3136	.2281	.2277
	O_LA	.0985	.0984	.0486	.0486	.0562	.0560
	R_SF	.0273	.0274	.0737	.0739	.0729	.0728
	R_SM	.2340	.2351	.1842	.1847	.2437	.2440
	R_LA	.1348	.1359	.0496	.0498	.0140	.0140

First column : predicted shares of housing alternatives under actual 1977 Federal Income Tax schedule (highest marginal tax rate: 70%).

Second column: predicted shares of housing alternatives under 1983 Federal Income Tax schedule, deflated by CPI to 1977 levels (highest marginal tax rate: 50%).

results in a slight shift from owning to renting. More comprehensively, the share of renting increases by:

Stratum	Albany	Dallas	Sacramento
Young Singles	.0005	.0004	.0017
Families	.0056	.0061	.0081
Eldery Couples	.0070	.0015	.0040
Widows	.0026	.0010	.0004

These numbers are very small: not only very few people are affected by the change in the marginal tax rate, but these "very rich" people are also those who are least likely to shift to the rental market.

Within each city, the shifts into rental units basically reflect the tax brackets which can be seen by a look at the mean income:

Stratum	Albany	Dallas	Sacramento
Young Singles	$ 5,200	$ 6,700	$ 5,200
Families	22,200	26,400	23,000
Eldery Couples	13,900	15,400	13,700
Widows	5,300	5,600	6,000

But mean income will not tell the entire story because the picture is complicated by distributional differences within each stratum and among SMSAs - both in terms of the income distributions and in terms of mortgage payments. This might explain the large shift to rental units among elderly couples in Albany.

Finally, the spill-over effects induced by the few moves in the rental market are calculated from the predicted moves in a stratum of

1000 nuclei (see the last column of Table 9). Note that the already mentioned problems with the small number of affected people are compounded by the erratic nature of the individual forecasts. The predicted changes in local property tax payments might therefore be unreliable. Aggregated over the three SMSAs and over all strata, the spill-over effect in lost property taxes is about 15 percent of the income tax deductions saved by the federal government. The latter are measured after the tax change: the percentage in terms of the direct effect is lower because the moves into the rental market partially offset the savings in income tax deductions.

We sum up the Federal Income Tax experiment as follows:

o Flattening the income tax schedule affects relatively few people and the changes in the aggregate are therefore small. Too few sample nuclei are affected to allow a reliable simulation.

o The pure price effect makes the fedaral income tax deductions worth less at high marginal tax rates. The resulting shift in the rental market is very small because the "very rich" people that are affected by the change are the least likely nuclei to switch to renting.

Caveats and Conclusions

Before drawing conclusions, two major conceptional caveats should be made: the first concerning the nucleus-approach, the other concerning the use of cross-sectional data.

To a good degree of approximation, households can be considered independent decision makers, i. e., for statistical purposes, the disturbances e_{it} and e_{js} are independent for different households $t \neq s$. But this does not hold for nuclei: e_{it} and e_{js} will be correlated if nucleus t lives in the same household as nucleus s. This intra-household correlation reflects the matching process, i.e., who shares a dwelling with whom, which is extremely difficult to model. As long as this matching process is uncorrelated with the disturbances in the housing choice, the intra-household correlation will only contribute to the variance in the estimates. But if the matching process and the housing choice have common unknown parameters, the estimates from our model are biased as well.

This statistical problem is substantially alleviated by stratifying the sample. The overwhelming majority of non-household heads are either adult children in their parents' homes or elderly parents in their childrens' homes. Thus, the correlated disturbances are in separate strata, and the aforementioned bias of standard errors or even parameters vanishes. The remaining multi-nucleus households that are not separated by differences in generation are roommates. Here it seems plausible to assume that their matching behavior is uncorrelated with their housing choices. For the purpose of this model, it can be assumed random. With the relatively low percentages of roommates among the non-heads (see Pitkin, 1980), the additional contribution to the variance is minor.

From the viewpoint of a model-builder, the more fundamental problem of the nucleus approach lies in the inefficient use of information rather than in the slidely biased variance. Information is lost by splitting up households into independent nuclei and separating them into different strata; e. g., it seems a valuable piece of information, whether an adult child has parents with a large house in town or not. In addition, the housing alternative "non-head" is a single category for a variety of rather different possible multi-nuclei households. As a special problem, the entire concept of headship is blurred in households of roommates, where no clear subordination exists.

The second caveat concerns the interpretation of the cross-sectional data as a steady state, especially with the prices as reported in the Annual Housing Survey. The approach ignores all intertemporal effects that might produce price dispersion or disquilibria. Spurious price elasticities may come from the fact that many sitting tenants receive tenure discounts: if we compare the rent of their actual unit with the hypothetical prices of those not chosen (measured as the prices paid by recent movers), the existence of tenure discounts will give us a larger price response than if we compare the prices with the tenure discounts subtracted. The same argument holds for other kinds of factors producing price dispersion in the housing market, e.g., search equilibria and explicit or implicit long run contract agreements.

There is empirical evidence for price dispersion: hedonic regressions for Albany and Dallas produce significant negative coefficients for length of tenure, indicating a 14 percent discount in Albany and a 4 percent discount in Dallas for a 10-year tenure (see Follain and Malpezzi, 1980). The failure of the compatibility test between static random utility maximization and the NMNL estimates in some of the married strata is a further hint that the true story might be intertemporal.

Both caveats have a common lesson: the model at hand reflects a reduced form or steady-state outcome of a mixture of rather complicated intertemporal processes, e.g., household formation, tenure discounts, search, and long term contracting. To identify the contribution of these processes to the steady state, longitudinal data are necessary. The panel from the last three waves of the Annual Housing Survey will be of special interest to the construction of structural models.

Taken as a descriptive device, the model performs well in terms of fit and prediction accuracy. Simulation results give a fairly stable pattern across SMSAs. In the case where the simulations coincide with other published experiments, the results were very close. All this gives us some confidence in the robustness of the model and its forecasts.

The main conclusion from the baseline estimates and from the housing allowance experiment is the strong response of headship rates to relative housing prices. Headship rates can not be treated as exogenous variables. The second conclusion concerns fiscal federalism: in all three fiscal changes, the spill-over effects from federal fiscal action to the local level and vice versa are of sizable magnitudes.

APPENDIX

NMNL-MODELS AND RANDOM UTILITY MAXIMIZATION

This appendix extends the Daly-Zachary theorem which provides the link between NMNL-models and the random utility maximization hypothesis (RUM).

Let c_T denote the similarity coefficient corresponding to the first-order clusters of elementary alternatives (say, tenure categories), and d_H the similarity coefficient corresponding to the second-order clusters consisting of first-order clusters (say, headship categories). The specified NMNL model is then equivalent to the following joint cumulative distribution function (McFadden 1978):

$$F(e_1,\ldots,e_M) = \exp \{ -G [\exp(-e_1),\ldots,\exp(-e_M)] \}$$

with

$$G[y_1,\ldots,y_M] = \sum_{\text{LIMBS}} \Bigl(\sum_{\text{BRANCHES}} \bigl(\sum_{\text{TWIGS}} y_i^{1/c} \bigr)^{c/d} \Bigr)^{d}$$

where we sum over the twigs (=elemental alternatives), the branches (=first-order clusters), and the limbs (=second-order clusters).

The connection between this c.d.f. and the corresponding density and the random utility hypothesis is given by the following two theorems:

Theorem 1 (Sufficiency) (McFadden 1979):

Let $0 < d_H < 1$ and $0 < c_T/d_H < 1$ for all T and H

Then the NMNL model is consistent with RUM for any data.

Theorem 2 (Necessity) (Daly and Zachary 1979):

Let $d_H > 1$ or $c_T/d_H > 1$ for at least one T or H

Then it is always possible to construct data at which the NMNL model is inconsistent with RUM. This failure occurs at points where the joint density f derived from F is negative.

The necessity argument by Daly and Zachary leaves the possibility open that for the data given by the application, the NMNL model is consistent with RUM, and that the data points where the inconsistency occurs are insensible for the given application. This gives rise to the question whether it is possible to construct a discrete choice model that is (1) compatible with RUM, (2) has the same cumulative distribution function F for the given data points, and (3) preserves the choice probabilities of the original NMNL model.

Proposition 1 (Sufficiency):

Let X be the set of all given data points. We assume X compact.

Let A be the open convex hull of X.

Let f denote the joint density function associated with F.

Let the following two conditions be met:

(1) f is non-negative in A,

(2) $M(A) := \int_A dF < 1.$

Then the data can be rationalized by a discrete choice model which is consistent with RUM and has the same cumulative distribution function F over A. However, this choice model will generally not have the same choice probabilities as the original NMNL model.

The idea underlying this proposition is to equally distribute the probability mass outside of A, $M(R^M-A)$, on the boundary of A, and redefine F outside the closed hull of A as zero. The equal distribution will generally distort the choice probabilities. A non-distorting distribution of $M(R^M-A)$ needs stronger conditions:

Proposition 2 (Choice Probability Preserving Choice Models):

Let B be the smallest open interval in R^M enclosing X.

Let N be the set $\{ x \text{ in } R^M \mid f(x) < 0 \}$.

Let P_i^{min} be the orthant-like support of $\inf \{ p_x(i) \mid x \text{ in } B \}$.

Let S_{ij} be the halfstrips between the sets P_i^{min}, such that R^M is partitioned into B, the P_i^{min}, and the S_{ij}.

Let L(y) denote the halfray defined by the origin y on the boundary of B and confined to the S_{ij} corresponding to y. For the corners of B, $L(y) := P_i^{min}$, i corresponding to y.

Let the following conditions be met:

(1) f is non-negative in the closed hull of B.

(2) $M(B) := \int_B dF < 1.$

In addition, let any one of the following conditions be true:

(3) N is a subset of one or more P_i^{min}.

(4) For any y on the boundary of B, $m(L(y)) := \int_{L(y)} dF > 0$.

Then and only then it is possible to construct a discrete choice model in B that (1) is compatible with RUM and has (2) the same cumulative distribution function and (3) the same choice probabililties in B as the original NMNL model.

The problem is depicted in Figure 3 for the three dimensional case with the normalized joint cumulative distribution function $F'(0,e_2-e_1,e_3-e_1)$. Condition (3) is based on the idea that shifting mass within the orthant-like sets P_i^{min} does not change any of the choice probabilities in B. If N has points outside the P_i^{min}, the negative mass can only be shifted along the halfrays L(y) without distorting choice probabilities inside B. Condition (4) ensures enough mass on each halfray to offset the points with negative density. Note that for L(y) at the corners of B, i. e., the P_i^{min}, condition (4) always holds. We now concentrate the mass $M(R^M-B)$ on the points y of the boundary of B in proportion to the L(y), and redefine F on R^M-B as zero. The so defined choice model has the claimed properties.

Finally, the necessity argument now follows as a corrollary:

<u>Proposition 3 (Necessity):</u>

Let any one of the following conditions be true:

(1) f is negative at a point in A.

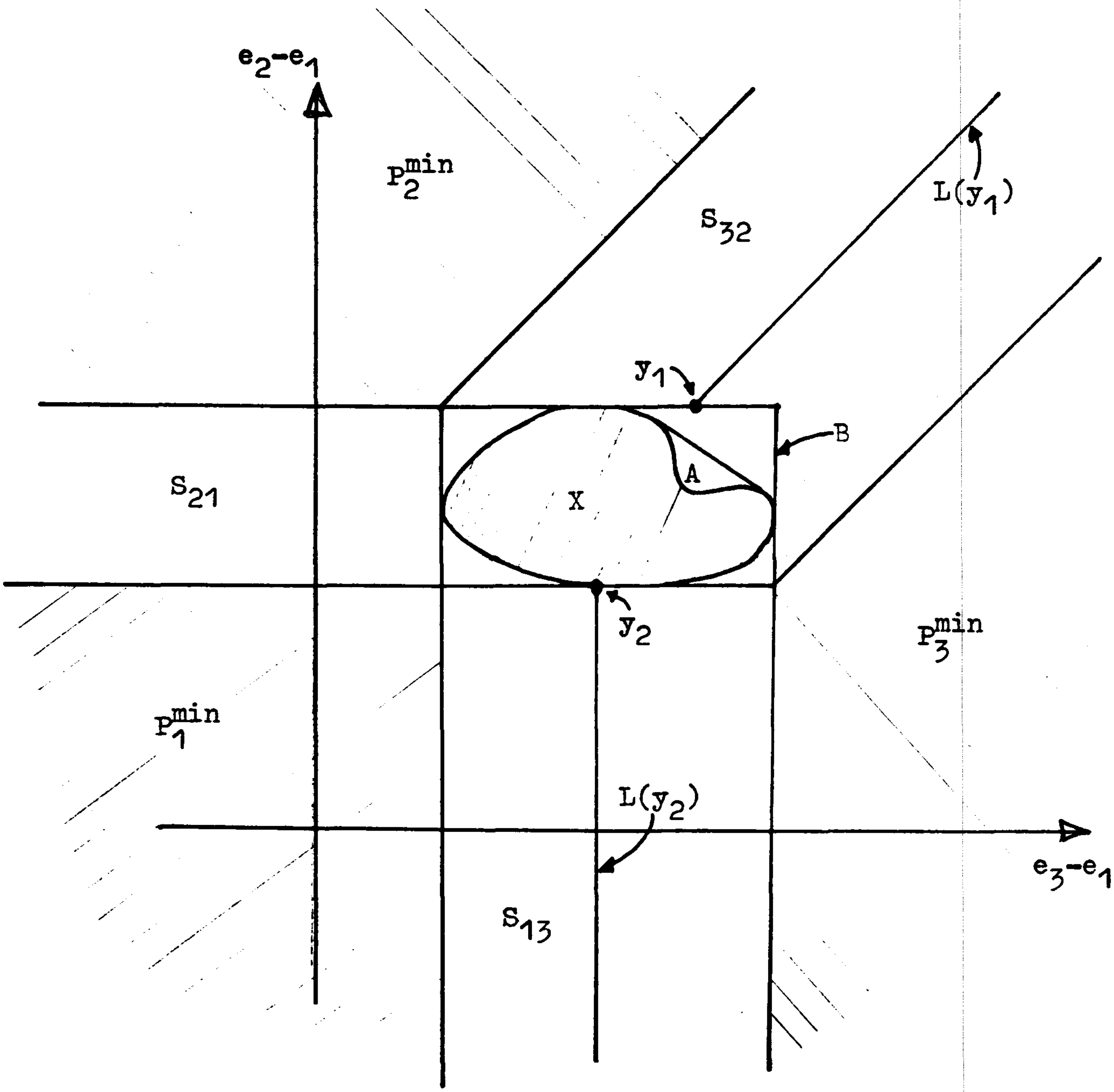


Figure 3

$$(2) \quad M(A) := \int_A dF > 1.$$

Then the construction of a RUM-compatible discrete choice model is not possible.

Proposition 1 through 3 exhaust all possibilities for X = B (=A). Exact proofs are tedious and can be obtained from the author at request. The first condition of Proposition 3 is violated by all four strata with similarity coefficients significantly larger than one. This renders these NMNL models irreparably inconsistent with the static microeconomic theory.

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