You Can Take It with You:

Proposition 13 Tax Benefits, Residential Mobility, and Willingness to Pay for Housing Amenities

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Abstract

In 1978, Californians approved Proposition 13, which fixed property tax rates at 1% of housing prices at the time of purchase. Beyond its fiscal consequences, Proposition 13 created a lock-in effect on housing choice because of the implicit tax break enjoyed by homeowners living in the same house for a long time. In this paper, I provide estimates of this lock-in effect, using a natural experiment created by two subsequent amendments to Proposition 13 - Propositions 60 and 90. These amendments allow households headed by an individual over the age of 55 to transfer the implicit tax benefit to a new home. I show that mobility rates of 55-year old homeowners are approximately 25% higher than those of 54 year olds. The second contribution of this paper is the incorporation of transaction costs, due to Proposition 13, into a household location decision model. The key insight of this model is that because of the property tax laws, different potential buyers have different user costs for the same house. The exogenous property tax component of this user cost then works as an instrument to solve the main identification problem of revealed preference models - the correlation between price and unobserved quality of the product. I find that marginal willingness to pay estimates for housing characteristics are approximately 100% upward biased when the model estimates do not account for the price endogeneity.

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Household sorting in the urban housing market has attracted the attention of economists since the pioneering work of Tiebout (1956).[1] Empirical research on local public finance, school choice, and segregation patterns, for example, have applied equilibrium sorting concepts originated in this literature.[2] In spite of its elegance, however, some of the Tiebout assumptions may not be credible, such as the free mobility of households. In reality, transaction costs and other barriers to sorting systematically affect individual behavior, although it is a difficult task to precisely measure those costs.[3]

In this paper I study the impact of one type of transaction costs – moving costs generated by property tax laws - on household mobility and how it can be used to recover preference parameters in a residential sorting model. The key insight is that in states where property taxes are based on historical prices rather than current market values, potential house buyers have *different* user costs for the *same* property. This research focuses on housing demand in California, where Proposition 13, passed in 1978, created unusually wide variation in property tax rates.

Proposition 13 replaced a decentralized system of property tax rates around 2-3% of assessed house values, with a uniform 1% fixed rate, based on prices at the time of purchase. The immediate effect of Proposition 13 was a one-time reduction in local property tax

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[1] Much of the intuition on household sorting was derived from a long line of theoretical work in local public finance that started in Tiebout (1956), and which includes Epple and Zelenitz (1981), Epple, Filimon, and Romer (1984, 1993), Benabou (1993), Nechyba (1997) and Epple and Sieg (1999).

[2] Recent examples are found in Barrow and Rouse (2000), Rothstein (2003) and Bayer, McMillan and Rueben (2002).

[3] As Rubinfeld (1987) points out "the value and usefulness of the Tiebout model is likely to diminish in the future, and an alternative or alternatives are needed." Also, see Quigley (2001) for a survey about the different types of housing transaction costs.

revenues.[4] The longer-run impact was to create a system of "grand-fathered" tax rates for houses based on historical prices. The associated tax savings can be substantial: Considering the one quarter of San Francisco Bay Area families with more than 20 years of housing tenure in 1990, I estimate that these savings amounted to an average of 4.5% of household gross annual income. The grand-fathering of tax rates therefore creates a "lock-in" effect, since a homeowner who moves to another home may experience a large increase in tax liability.[5]

However, under a pair of propositions passed in the late 1980's (Proposition 60 in 1986, and Proposition 90 in 1988), homeowners aged 55 or older who sell a property and buy another of equal or lesser value are allowed to keep the tax base value of their original home. These laws created a sharp discontinuity in the lock-in effect of Proposition 13, giving rise to an interesting natural experiment for estimating the impact of moving costs on mobility.

I estimate the lock-in effect attributable to Proposition 13 by comparing householders who are 54 years old to those who are 55. I find that 55-year olds have a 1.2-1.5 percentage point higher rate of moving (on a base of approximately 4%). Consistent with a tax-based explanation for this difference, 55-year old recent movers paid 15% less property taxes than their 54-year old counterparts. To check whether this change in mobility is due to other discontinuous trends, I look at mobility rates for various control groups, including California homeowners in 1980 and renters in 1990, and Texas homeowners in 1990. In all, I find no evidence of a discontinuity. Moreover, there are no differences in property taxes paid by 54 and 55-year old recent movers for these control groups.

Transaction costs due to Proposition 13 are then explicitly incorporated in a household

[4] Total property tax revenues in California declined by 45% in 1978-1979, and the share of local counties` revenue from property taxes declined from 33% in 1977-1978 to 11.6% in 1995-1996 – see Shapiro and Sonstelie (1982), Rosen (1982), Silva and Sonstelie (1995), Fischel (1989), and Brunner and Rueben (2001) for fiscal consequences of Proposition 13.

[5] These calculations are explained in Section 3. The Proposition 13 effect is analogous to the spatial lock-in related to falling housing prices, as in Caplin, Freeman and Tracy (1997) and Chan (2001), or due to increase in interest rates, as in Quigley (1987).

location decision model. The output from this revealed preference model consists of a set of underlying taste parameters for housing and neighborhood characteristics, which are of special interest for understanding sorting patterns and valuation of local public amenities. Here I adopt estimation strategies first used by McFadden (1974 and 1978), and updated by Berry (1994), Berry, Levinsohn and Pakes (1995), and Petrin and Train (2004).

There are two main differences between the household sorting model developed in this paper other revealed preference models. First, instead of assuming that consumers face one *fixed* price for each house, Proposition 13 and its amendments naturally generate a *specific* user cost of the house to each homeowner. Second, I use the variation in moving costs created by Propositions 13, 60 and 90 as an instrument to control for the correlation between price and the unobserved housing quality. This is the first time one uses such research design, which incorporates a reduced form identification strategy in a structural model of the housing market, in order to account for the endogeneity of housing prices. [6]

The implementation of this sorting model is only feasible using the 1990 California Decennial Census Long Form data, which is a 15% sample. These are restricted-access micro data, with information for approximately two million households in California, including the property taxes paid by each. Unlike the publicly available micro sample, in which the smallest geographic area contains 100,000 individuals, the 15% sample reveals the location of each house and work place at the Census block level, a region with approximately 100 individuals. This special feature allows me to precisely define neighborhoods, and at the same time incorporate a rich set of observed heterogeneity, such as income, race, age and distance to work.

Simple multinomial logit estimates of the sorting model generate a relatively small user

[6] Credible instruments to control for the correlation between prices and unobserved housing quality are scarce in this literature. Bajari and Kahn (2001) estimate bounds on willingness to pay for distance to work in order to avoid the use of instruments. Bayer, Ferreira and McMillan (2003) instrument price with a quasi optimal instrument derived from the choice model and from land use measures, to estimate valuation of school quality.

cost coefficient, indicating very high preferences for certain housing characteristics. This result is typical of an omitted variable bias situation: given that we do not observe all housing amenities, prices tend to be higher for houses with valuable unobserved attributes. This problem can be solved by including a control function, in which differences in the tax cost across houses attributable to Proposition 13 work as an instrumental variable for the user cost of the house.

Preference parameter estimates from the adjusted model are then used to recover estimates of the marginal willingness to pay (MWTP) for housing and neighborhood attributes. I find that MWTP estimates for housing characteristics are 100% upward biased when not controlling for the endogeneity of housing prices. For example, homeowners are willing to pay, on average, annually \$2,000 for one extra room and \$4,450 for a detached house, compared to \$3,994 and \$9,224 of a simple logit model. These results hold after the inclusion of heterogeneity and wealth effects. Interestingly, the same estimation method breaks down when applied to Texas, given the lack of meaningful variation in property taxes for that state.

The rest of this paper is organized as follows. In Section I, I explain Proposition 13 in detail. In Section II, I describe the data set, while in Section III, I present estimates of the lock-in effect. Section IV presents a household residential location model, and Section V provides estimates of MWTP for housing characteristics. Section VI concludes.

I. Proposition 13

Proposition 13 was approved in 1978 by 65% of the voters in California. The vote was widely interpreted as a "tax revolt" against the state government. In the mid-1970's, property tax revenues in California were quickly fueled by sky-rocketing house prices and the unwillingness of local officials to cut property tax rates in the face of a growing tax base.

Advocates of the proposition argued that tax increases were forcing elderly and low-income families to sell their homes. At the same time, school spending in the state was dramatically changing in response to the California Supreme Court's decision, *Serrano vs Priest* (1971), which required the equalization of spending per pupil across school districts. Fischel (1985) argues that the cost of the equalization program provoked a reaction by the voters in the form of restricting government revenues through Proposition 13.[7]

According to the California Constitution, Article XIIIa, Proposition 13 states that "the maximum amount of any ad valorem tax on real property shall not exceed one percent (1%) of the full cash value of such property". Full cash value means price at the time of purchase plus a maximum inflation adjustment of 2 percent per year. No re-assessment could be carried out, implying that property taxes are effectively frozen (apart from the 2% per year rise). Also, the initial base values used to set property taxes were the assessed housing values of 1975/1976.[8]

Two important modifications to Proposition 13 were enacted during the next decade. Proposition 60 was a constitutional amendment approved in 1986, which allowed the transfer of tax benefits for within-county movers. Proposition 60 permits a transfer of a Proposition 13 base year value of the property from the current residence to a replacement dwelling if: a) homeowners are at least 55 years old; and b) the replacement dwelling is of equal or lesser value than the selling price of the old property. In practice, Proposition 60 enabled 55-year or older households to carry the frozen property taxes to a new home within the same county.

Proposition 90, approved in 1988, brought even more flexibility, allowing inter-county base year value transfers. Adoption of Proposition 90 was not mandatory and the law only

[7] Proposition 13' effects still reverberate today, as recurrent state budget pressures lead to under funding of several public services. Paul Krugman wrote in New York Times editorial of 08/22/2003 "Thanks to Proposition 13, some people pay ridiculously low property taxes. Warren Buffett offered the perfect example: he pays $14,401 in property taxes on his $500,000 home in Omaha, but only $2,264 on his $4 million home in Orange County."

[8] The law also says that the limitation of 1% of property taxes for all local governances would not apply to additional taxes to pay for specific bonds approved by a super-majority of voters. Also, Proposition 13 required that any new taxes proposed by the state legislature had to be approved by a two-thirds majority of each house.

applies across counties that approved the ordinance. Only a few, albeit relatively large, counties in California adopted Proposition 90 immediately after approval of the law, namely: Alameda, Contra Costa, Inyo, Kern, Los Angeles, Marin, Modoc, Monterrey, Orange, Riverside, San Diego, San Mateo, Santa Clara, and Ventura.[9]

II. Data Set

I estimate the lock-in effect of Proposition 13 using the Integrated Public Use Microdata Series (IPUMS) 5% samples of the 1980 and 1990 for the states of California and Texas. The choice of Texas as comparison group comes from the fact that house values are re-assessed every two to three years in that state. Table 1 shows average house values, property taxes, and effective tax rates (individual house values divided by property taxes) for the full sample, and by the date when households moved into their home.

A striking feature of these data is the gap in effective property taxes paid by homeowners of different tenures in California in 1990. While homeowners who had moved in the previous year paid an effective tax rate of 0.8% on average, households living in the same dwelling for more than a decade paid less than 0.44%. This discrepancy corresponds to a tax saving of $900 per year in 1990 dollars. If we focus on the implicit tax benefit – the difference between current property taxes and 1% of house values - for households who moved in before 1979, this number can reach almost 3% of household gross annual income. In some places, such as the San Francisco Bay Area, the implicit tax benefit reached almost 4.5% of gross income for the same selected group of households.

When looking at California in 1980, by comparison, we only see a small difference in effective property taxes between homeowners who moved before 1975 and those who moved

[9] 60% of the state population is located in these counties. Four of those counties have subsequently repealed the ordinance: Contra Costa, Inyo, Marin and Riverside.

after. This is the initial consequence of Proposition 13, when property taxes were set at 1% of house values assessed in 1975. As opposed to California, the Texas data show relatively stable effective property tax rates. Only households who moved before 1970 have discounts in property taxes, presumably because Texas offers special exemptions for householders 65 years of age and older.[10]

Figure 1 plots effective property tax rates by age for California homeowners in 1990. The distributional effects of Proposition 13 are clear: elderly households pay less property tax than younger households, because of differences in tenure by age. When normalizing property taxes by annual household income instead of house values (Figure 2) the distributional effects of Proposition 13 are less pronounced. The main characteristic is that individuals between 50-60 years of age pay less tax as a proportion of their income compared to other age groups. This might reflect the age profile of income, where maximum income is generally achieved around age 50. In comparing current taxes with a counterfactual 1% of housing values as property taxes, the gap between what Californians should pay in a different regime is much larger for the elderly.

III. Lock-in Effect

The lock-in effect of Proposition 13 arises because of the implicit tax break for households who have been living in the same house for a long time. This paper is the first research to identify the lock-in, by looking after age 55, when the lock-in effect is removed.[11] Figure 3 illustrates the key insight of the research design. It graphs the probability of moving to a new house in 1990 by age group. Each dot in Figure 3 is calculated as the total number of

[10] Appendix Table 1 reports descriptive statistics for housing and homeowner characteristics.

[11] Few papers study the effects of Proposition 13 on mobility, such as Sexton, Sheffrin and O'Sullivan (1995) and (1999). Nagy (1997) looks at mobility rates before and after the law approval, finding no significant effects (the lock-in effect could only have an impact after a significant house prices increase).

homeowners who moved in the last year, divided by the total number of homeowners from the respective age. From now on, age is defined as the maximum age between householder and spouse, to correspond with the provisions of Propositions 60 and 90.

A sharp discontinuity arises between 54 and 55-year olds. The probability of moving for a 54-year old is 4% while for 55 year olds it reaches 5.2%. This 1.2% point difference is presumably caused by the effect of Propositions 60 and 90. The remainder of this section presents a variety of tests of this interpretation.

In order to rule out competing hypothesis, I compare 1990 California data with several control groups, such as California data from 1980, before Propositions 60 and 90 had been approved. Figure 4 graphs the probability of moving for this group, where only the negative relationship between mobility and age is found. This comparison rules out any type of special Californian mobility pattern as the explanation for the sharp change in mobility rates.

Figure 5 plots the probability of moving for renters in California 1990. Again, no discontinuity is found for the relevant age group. The existence of a 1989-1990 localized year effect on mobility is ruled out by this comparison. Figure 6 plots the probability of moving for homeowners in Texas in 1990. Again, no discontinuity is found, allowing me to rule out national economic shocks, regulations or trends as cause of the change in mobility rates for 55 years old in California in 1990.[12]

Table 2 reports results from a probit model, designed to quantify the patterns observed in the figures above. The following reduced form equation for the probability of moving in 1990 is estimated:

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[12] It is important to compare mobility rates with other states because the old capital gain laws also allowed homeowners 55-year or older to avoid certain payments under some conditions. When comparing mobility rates for the state of Massachusetts, where house prices significantly increased in the 1980's, there is also no differences in mobility rates between 54 and 55-year old homeowners. For the effect of other tax reforms on housing demand see Sinai (1997).

$$\text{(1)} \qquad \Pr(moving)_i = \Phi(\delta_1 D_i^{55} + \delta_2 Age_i + \delta_3 Age_i^2 + \delta_4 Age_i^3)$$

where D_i^{55} is a dummy for 55-year or older and $\Phi(\cdot)$ is the normal c.d.f.. The age controls are included in the equation because the effect of age on mobility is non-linear. Column (1) shows a negative correlation between D_i^{55} and the probability of moving, due to the negative impact of age on mobility rates. Column (3) adds the polynomial in age, leading to a change in the sign of the age 55 and older dummy, and setting the effect of D_i^{55} on mobility in 1.5% points. The estimated parameter is unchanged with the addition of housing attributes, household characteristics or fixed effects at the metropolitan area - this result is typical of a regression discontinuity design: covariates should not matter around the threshold that defines a treatment and control groups.[13] Pooling the 1990 California data with 1980 California data or the 1990 Texas data increases the estimated effect to 1.7% and 2% respectively, which is consistent with the downward trend in mobility rates observed in those control groups. Finally, I exclude 54 to 55-year households from the sample, in order to control for possible measurement errors in age. This test also verifies the existence of an overall effect in mobility patterns as opposed to only 54-years homeowners delaying mobility until they are 55 years old. The estimate of 1.4% confirms that Propositions 60 and 90 affected all homeowners older than 55.[14]

It is important to note that the reduced form results hold for the full local population of 54 and 55-year old, independent of their moving status. Given the 1-year difference in both

[13] See Cook and Campbell (1979).

[14] A remaining question relates to how permanent or transitory were the effects of Propositions 60 and 90. Given that we are looking at mobility rates in 1989-1990, 3 years after Proposition 60's approval and 1 year after Proposition 90, potentially these analyses capture mobility for a stock of households that were mismatched for some period of time and not only in the year period. Although this is a relevant consideration, there is no evidence that homeowners were moving at higher rates to houses located in different counties because of Proposition 90 (which was approved only 1 year before the period of analysis).

cohorts, there is no reason to expect differences in preferences or average characteristics of those households. Figure 7 confirms this result by plotting a set of house and neighborhood characteristics by age, which shows no evidence of a discontinuity between 54 and 55-year olds. All comparisons above point out to a causal relationship between the ability to transfer the tax benefit and mobility rates.

A. Robustness Checks

The main consistency check relates to the ability of transferring the tax benefit. If recent movers in fact used Propositions 60 and 90, a discontinuity in property taxes payments would be expected. Figure 8 shows average property taxes by age. The gap between 54 and 55 year olds is approximately $220 per year, and is only noticeable in California in 1990. Only a downward trend in property taxes payments is observed in California in 1980, and Texas in 1990. It is worth noting that in Texas, 65 year of age or older households enjoy several exemptions in their tax payments. Table 3 reports differences in effective property tax data , where differences are estimated for recent movers and long tenure homeowners. Those differences are only significant for recent movers in California, 1990.[15]

The $220 per year gap in taxes between 54 and 55 year old Californians in 1990 seems a small number compared to the differences in property tax payments reported in Table 1. If long tenure households were moving in 1989-1990 in similar proportions, i.e., "when moved in" groups were contributing with proportional number of recent movers, the expected average gap

[15] Differences in property taxes and effective tax rates in California in 1990 can only be used as a consistency check: given that Figure 8 and Table 3do not report results for the overall population, any comparison between 54 and 55-year olds would suffer from selection bias.

would be $536 per year.[16] This indicates that long tenure homeowners were probably moving with lower rates than short tenure homeowners. Also, 55-year old homeowners moving to more expensive houses are not allowed to transfer the tax benefit.

Families moving to counties that did not allow Proposition 90 are another explanation for the modest tax difference. Figure 9 shows the probability of moving for California 1990 split in two groups: movers who could transfer the tax benefit (because of Proposition 60 or 90) and movers who could not (because Proposition 90 was not allowed). The comparison is made using the Census question: "Where did this person live 5 years ago (on April 1, 1985)?" 22% of the 54 and 55-year olds recent homeowners moved to places that did not accept Proposition 90. Figure 10 also allows me to calculate the discontinuity in both groups. Only the group allowed to transfer the tax benefit had a gap of .95% points between probabilities of moving for 54 and 55 years old. The same comparison is made in Figure 15, but plotting average property taxes instead of mobility rates. Not surprisingly, the gap between 54 and 55-year old increased when comparing the predicted average property taxes.[17]

IV. Residential Location Decisions Model

In this section I develop a household residential demand model, where the focus is on the incorporation of transaction costs represented by Proposition 13. Taxation costs are included into the model and used as a device to recover estimates of the marginal willingness to pay (MWTP) estimates for housing and neighborhood attributes. The key insight of the model is the construction of a user cost for a house that varies across people. The property tax

[16] In order to estimate the full tax benefit, there would be the need to know the expected tenure for each homeowner, compare it with their expected future income (which is potentially decreasing given the proximity to retirement age), and then calculate present values.

[17] A final explanation is that some of the new movers may have been renters in the previous house. For example, the proportion of 54-55 years old non-movers who are renters is 20% for California in 1990. Also, from the March CPS question "What was (your/name) main reason for moving?" we can infer that 16.2% of the 50-59 years of age households pointed out "wanted to own home, not rent" as the main reason to move.

differences created by Proposition 13 provides exogenous variation in user costs, and can be used as an instrumental variable to reduce the influence of unobserved house characteristics in estimating MWTP.

The model is based on standard differentiated product demand models, whose roots lie in the work of McFadden (1973,1978) and more recently Berry (1994), Berry, Levinsohn and Pakes (1995), and Petrin and Train (2004).[18] The central idea is that demand parameters can be recovered from observed choices in the housing market, where houses are considered as bundles of characteristics. Households choose to live in the house that maximizes expected utility derived from housing and location attributes.

A number of existing studies have used similar or related frameworks to estimate preferences for housing and neighborhood characteristics. Palmquist (1984) directly estimated demand for certain house characteristics in seven metropolitan areas using the hedonic approach developed by Rosen (1974). Recently, several papers adapted the BLP approach to the housing market, including Bajari and Kahn (2000), Bayer, McMillan and Rueben (2003) and Bayer, Ferreira and McMillan (2003). These last two papers develop an equilibrium model of the housing market, allowing the estimation of general equilibrium simulations to evaluate changes in policy. None of these papers, however, explicitly takes into account the variation in user costs of alternative housing units posed by Proposition 13 or similar laws in other states.

A. The Model

Assume that household i maximizes utility by choosing among alternative houses indexed by j. Also, assume that housing supply is fixed. The indirect utility of household i from

[18] For a detailed explanation of the random coefficients multinomial logit model, see Nevo (2000). A review of the earlier literature can be found in Train (2000).

consuming house *j*, $U(p_j, \tau_{ij}, x_j, z_i, \xi_j; \theta)$, is defined as a function of housing prices p_j, the property taxes paid by each homeowner τ_{ij}, a vector of housing amenities x_j, a vector of observed household characteristics z_i - including annual household income I_i, unobserved attributes of the house ξ_j and a vector of unknown parameters θ defining mean and heterogeneity in preferences. I adopt the following functional form:

(2) $$u_{ij} = \alpha_i g(I_i - p_{ij}) + x_j \beta_i + \xi_j + \varepsilon_{ij}$$

where $g(\cdot)$ is a monotonic function, ε_{ij} is the stochastic term, and α_i and β_i are preferences for housing prices and attributes. Each parameter associated with the choice variables in the model varies with a household's own characteristics according to:

(3a) $$\alpha_i = \alpha_0 + \sum_{r=1}^{R} \alpha_r^m z_{ri}$$

(3b) $$\beta_i = \beta_0 + \sum_{r=1}^{R} \beta_r^m z_{ri}$$

and equations (3a) and (3b) describe household *i*'s preference for housing characteristic *m*.[19] The term p_{ij}, which I call the user cost of the house in a given year, is defined as:

(4) $$p_{ij} = r p_j + \tau_{ij}$$

[19] Unobserved heterogeneity is not modeled in this paper because of two reasons. First, the microdata allow me to incorporate a rich set of observed heterogeneity that gives rise to flexible substitution patterns. Second, it is still part of future research how to incorporate unobserved heterogeneity in models that use random samples of alternatives as a choice set.

where r is the annual interest rate. The user cost of the house is composed by a common carrying cost rp_j faced by all individuals, and property taxes τ_{ij} specific to each homeowner.

Alternative choices for the function $g(\cdot)$ determine whether there are income effects in the marginal willingness to pay for amenities. The MWTP by household i for amenity j is:

$$(5) \qquad MWTP_{ij} \equiv -\frac{\dfrac{\partial u_{ij}}{\partial x_j^m}}{\dfrac{\partial u_{ij}}{\partial p_{ij}}} = \frac{\beta_i^m}{\alpha_i} \frac{1}{g'(I_i - p_{ij})}$$

If $g(I_i - p_{ij}) = I_i - p_{ij}$ then the MWTP is just $-\beta_i^m / \alpha_i$. On the other hand, if $g(I_i - p_{ij}) = \log(I_i - p_{ij})$, as would be the case under a Cobb-Douglas specification of preferences, then:

$$(6) \qquad MWTP_{ij} = \frac{\beta_i^m}{\alpha_i}(I_i - p_{ij})$$

which is increasing with income net of housing costs.

Given the household's problem described in equations (2)-(4), household i chooses housing choice j if the utility that it receives from this choice exceeds the utility that it receives from all other possible house choices, i.e.,

$$(7) \qquad u_{ij} > u_{ik} \quad \Rightarrow \quad W_{ij} + \varepsilon_{ij} > W_{ik} + \varepsilon_{ik} \quad \Rightarrow \quad \varepsilon_{ij} - \varepsilon_{ik} > W_{ik} - W_{ij} \qquad \forall \quad k \neq j$$

where W_{ij} includes all of the non-idiosyncratic components of the indirect utility described in (2). As the inequalities in (7) imply, the probability that a household chooses any particular choice depends in general on the characteristics of the full set of possible house choices.

Assuming ε_{ij} follows an iid extreme value distribution, the probability of household *i* choosing house *j* from choice set *J* has the following functional form:

(8) $$\Pi_{ij} = \frac{\exp(\alpha_i g(I_i - p_{ij}) + x_j \beta_i + \xi_j)}{\sum_{j=1}^{J} \exp(\alpha_i g(I_i - p_{ij}) + x_j \beta_i + \xi_j)}$$

Maximizing the probability that each household makes the correct housing choice gives rise to the following log-likelihood function:

(9) $$L = \sum_i \sum_j 1_{ij} \ln(\Pi_{ij})$$

where 1_{ij} is an indicator variable that is equal to one if household *i* chooses house *j* and zero otherwise.

B. Endogeneity of Housing Prices

The main concern that arises is estimating MWTP in the framework of equations (2)-(9) comes from the correlation between price and the unobserved portion of the utility. This correlation is caused by omitted variables – the econometrician does not observe all characteristics of the house that affects utility, i.e., prices tend to be higher for houses with valuable unobserved attributes.

Most papers on demand for differentiated products have used two methods to solve this problem: the control function approach or the Barry, Levinshon and Pakes (1995) method.[20] The main problem with both approaches is the difficulty in finding instruments correlated with price and uncorrelated with the mean utility that all households share from each house. In this paper, I choose the control function approach because of the high number of products in the housing market, leading to complications in estimating mean utilities for each product in a Barry, Levinshon and Pakes (1995) framework.[21,]

The variation in taxation costs faced by homeowners in California is key element in the identification strategy. The instrument corresponds to the "clean" variation in p_{ij}, i.e., corresponds to the component of the user cost of the house dependent on Propositions 13, 60 and 90. Proposition 13 gives the variation in implicit tax benefits faced by homeowners, while Propositions 60 and 90 set the moving costs when householders decide to choose another property.

A potential concern about using moving costs as source of identification is the instrument's correlation with homeowner tenure. If the instrument is correlated with tenure, then it is also potentially correlated with unobserved housing quality, undermining the identification strategy. A simple regression of property taxes on tenure produces an R^2 of .15. In order to address this issue, an alternative specification is used for the instrument, where

[20] In the control function, a set of instrumental variables is used in a first stage regression of prices on housing attributes. In the second stage, a function of the first stage predicted residuals is included in the choice model. In the Barry, Levinshon and Pakes (1995), a series of mean utilities derived from market shares are estimated in the choice model. The mean utilities are then regressed on price, housing variables, and the price instrument.

[21] Hausman (1978), Heckman (1978) and Smith and Blundell (1986) initially developed the control function, and it can be thought as a two-stage least square approach applied to non-linear models. Blundell and Powell (2001) expanded the control function ideas in a semi-parametric and non-parametric estimation. It is important to emphasize that the control function approach does not have the same properties of the traditional Barry, Levihson and Pakes (1995) approach. Petrin and Train (2004) demonstrate the differences between the two methods.

individual tenure is decomposed out of moving costs.

The procedure works as follows: First, the effective property tax rates are estimated for each homeowner. Then, block group average effective tax rates are calculated for homeowners with age groups, such as 30-34, 35-39, 40-44 year olds, and so on. Finally, the house value of each homeowner is multiplied by the relevant average effective tax rate. In doing so, the correlation of the instrument with individual homeowner tenure is mitigated, leaving an adjusted property tax that is a function of past housing values in the same neighborhood for people of the same age group. As expected, I find an R^2 of only .04 for the regression of the adjusted instrument on homeowner tenure. I use this adjusted instrument throughout the paper.

In practice, I estimate the following first stage for the income effects specification:

$$g(I_i - p_{ij}) = \lambda \tau_{ij} + x_j \psi + v_{ij} \tag{10}$$

Then, the predicted residual $\hat{v}_{ij}$ is incorporated in the utility function as a linear term:

$$u_{ij} = \alpha_i g(I_i - p_{ij}) + x_j \beta_i + \delta_i \hat{v}_{ij} + \varepsilon_{ij} \tag{11}$$

where δ_i also depends on observed household characteristics. As evident from equation (11), the predicted residual $\hat{v}_{ij}$ is a proxy for the unobserved housing quality ξ_j.[22] As in traditional two-stage least squares estimates, the identification strategy fundamentally relies on the first stage results.

[22] I only include a linear function of the residual in the estimates, although the control function allows the inclusion of any non-linear function. As a consistency check, interactions of the predicted residual with choice variables were included in the model, leading to results relatively similar to the original specification.

While price endogeneity is the main identification problem of revealed preference models, it is not the only one. In order to estimate the model, it is assumed that house characteristics are uncorrelated with the unobserved portion of the utility. As an example, house style or front yard size are assumed to be uncorrelated with number of rooms. If this is not the case, MWTP estimate for an extra room will be biased. Although this might seem a restrictive hypothesis, to my knowledge there is no paper in the housing literature that addresses this question by using exogenous variation from a natural experiment.

V. Estimation Results

This section presents estimation results for the preceding model, using data on 98,407 homeowners between 30 and 70 years of age living in the San Francisco Bay Area and included in the 15% restricted use 1990 Census sample[23]. I use the restricted California Decennial Census Long Form data because in addition to containing the location of each house at the block level, the Long Form database also includes more complete data on key variables, such as property taxes. In particular, although the public use files of the Census top code property taxes at $5,000 and report only discrete ranges of taxes, the restricted Long Form data have the exact property tax paid by all households up to a $15,000 cap.

The analysis is restricted to residents of a single metropolitan area for several reasons. First, it is a self-contained economic region, with small proportion of commuters in and out of the region. Second, by focusing on a single metropolitan area, I restrict attention to alternative housing choices in the same area. Finally, for reasons of tractability and for obtaining permission to use the restricted Census data it is more convenient to use data from a single metropolitan area.

[23] The sample is composed of six counties: Alameda, Contra Costa, Marin, San Jose, Santa Clara and San Francisco.

In the estimation, each household is assumed to compare the value of their current house to the value of a set of alternative houses. I assume that the set of possible alternatives includes houses that were newly purchased in the previous year. This is best proxy for houses available in the market in the year of analysis.[24] For each household in the estimation sample, I randomly assign 10 alternative houses from the choice set - the consistency of this procedure is guaranteed by the IIA property, as demonstrated in McFadden (1978).[25]

The choice variables include characteristics of the house (draw from the Census data), socio-demographic characteristics of the neighborhood (based on averages at the block group level from the Census data), and characteristics of the neighborhood from external data, including elevation, population density, a measure of local air quality and a measure of 1st grade test scores in the nearest public primary school.[26] Appendix Table 2 shows the average characteristics of the houses owned by people in the sample and of the alternative houses. The alternative houses have a smaller number of rooms, were built more recently and are more likely to be apartments or attached dwellings. Neighborhood characteristics are very similar for both groups, although chosen houses are located in slightly whiter and richer block groups.

From the point of view of the model developed in the last section, the most interesting feature of the chosen house versus the alternatives is the property tax. For the chosen houses,

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[24] Misspecification of the choice set may lead to serious estimation biases. Swait (1984) showed, for example, that not incorporating captivity to a certain group of alternatives, lead to downward biased estimates for choice characteristics and upward biased fixed effects parameters. The logic is simple: when we include in the model alternatives not available to individuals (or not considered by them), we are in fact adding extra noise, which will be captured by the fixed effects, reducing the importance of observed choice variables.

[25] Although IIA property dictates substitution patterns among individual alternatives, the inclusion of observed heterogeneity allows flexible substitution patters at higher levels of aggregation. Also, the inclusion of distance to work gives rise to more reasonable substitution patterns in the urban space.

[26] Elevation is measured at the block level (source: EPA: BASINS - Better Assessment Science Integrating Point and Nonpoint Sources). Population density combines Census data and block group areas drawn from ArcView GIS. Average test scores of 1991-1992 and 1992-1993 academic years are assigned from the closest school within the school district, using census block centroids and school latitudes and longitudes (source: California Department of Education, 1991-1993). Air quality is predicted for each census block using information from monitor stations (source: Rand California, 1990) and industrial plants (source: EPA – AIRS –Aerometric Information Retrieval System).

the property tax is reported in the Census. For the alternative, the institutional framework of Propositions 13, 60 and 90 is used to generate the taxation costs that a specific household faces when choosing that house. For example, a 30 year old choosing a house from the alternative set is assumed to have property taxes calculated as 1% of the house value.[27] On the other hand, homeowners age 55 or older are allowed to transfer current property tax cost of their current home to another house if: a) the housing alternative is of equal or lesser value; b) the homeowner is moving within the same county or to a county that accepts Proposition 90. After generating property taxes for all alternatives, the individual user cost of the house is constructed as in equation 8, using an interest rate of 6%.[28] Appendix Table 3 reports the first stage estimates of the user cost on property taxes and housing and neighborhood variables for the pooled set of houses and alternatives. As expected, all specifications show a high F-test for the instrumental variable.

Multinomial logit estimates are presented in Table 4. Column (1) shows preference parameters for a model without heterogeneity and assuming that utility is linear in income net of housing costs. I focus on three variables - number of rooms, detached houses and average income of the neighborhood – to compare how changes in the model affect housing and neighborhood MWTP estimates. All signs look correct – negative for price and positive for the choice variables. The main problem is the magnitude of the price coefficient, which suggests a very small value for the marginal utility of income, or alternatively very high value of willingness to pay. This problem, which has been noted in other studies, is arguably due to the fact that house prices are correlated with unobserved characteristics of the house. Looking at columns (2)-(4), the estimated coefficient of the user cost variable remains relatively small in magnitude, even when including a broad set of housing and neighborhood controls.

[27] Median self-reported house values from the Census are used to calculate this cost.
[28] Similar results were achieved with interest rates of 8% and 10%.

Column (5) reports the estimate results for a specification similar to the one in column (4) but with the addition of a control function, equal to the residual of the first stage model for the user costs, as indicated in equations (10) and (11). The coefficient on the control function is large and positive, suggesting that unobserved variables that affect price also affect the utility assigned to the house. When the control function is included, the coefficient on the user cost rises in magnitude by a factor of 3.

MWTP estimates derived from the primitives of the model are shown in Table 5. The simplest model shows very high MWTP estimates for housing characteristics - \$3,275 per year for an extra room and \$17,210 for a detached house. On the other side, MWTP for average income of the neighborhood seems too small - \$1,081 for a \$10,000 higher average income of the neighborhood. Even after including other controls for housing and neighborhood amenities, the results still look very similar. Column (5) shows the inclusion of the control function. As noted in the multinomial logit estimates, the control function has the expected effect of deflating the MWTP estimates for housing characteristics - \$2,021 for number of rooms and \$4,446 for a detached house.[29]

Wealth effects are included in columns (6) and (7) of Tables 4 and 5. Again, the results are meaningful only after controlling for the unobserved housing quality. The user cost coefficient sign is opposed to the initial estimates because it was replaced by the income net of housing costs. The MWTP estimates are \$2,083 for number of rooms, \$6,706 for a detached house and \$1,616 for average income of the neighborhood. When wealth effects from owning a property are considered in the model, homeowners are on average willing to pay a similar amount for an extra room, but 50% more for a detached house and 25% more for a

[29] Appendix Table 4 reports results for a standard hedonic regression model. Although the hedonic model is not purged of any of the selection and endogeneity problems, it offers a good benchmark for comparing the model estimates.

neighborhood with higher average income.

Finally, on Table 6 I include observed heterogeneity in the model. Household income, age, and a dummy for white are interacted with all choice characteristics, including distance to work. Column 1 shows baseline MWTP estimates, corresponding to a mixed race household (given by the share of whites in the sample) with average income and average age. Although not directly comparable, those MWTP are strikingly similar to the model that assumes homogenous preferences. Columns (2), (3) and (4) report MWTP results from a household with higher income, higher age, and white respectively. The $10,000 change in income is not enough to have a noticeable effect in preferences. 10 years of age has a sizable effect though, especially on the choice of housing type. Finally, white households have strong preferences to live with individuals of the same type.

A. Robustness Checks

In order to assure that estimation results are driven by the variation in the data, and not by the structure of the model, I also estimate a similar housing choice model for Dallas-TX in 1990.[30] Table 7 shows the multinomial logit estimates for models with and without the control function. The standard multinomial logit showed in column (1) has a small and negative user cost coefficient, as we observed for the California estimates. However, the results are very different for the control function specification showed in column (2). The control function coefficient estimate is not significantly different than zero, and its magnitude is very small and not powerful enough to shift the user cost coefficient. MWTP estimates are reported in columns (3) and (4), and they show very similar results for both approaches, indicating that the

[30] Homeowners age 65 or older were excluded from the sample because they receive several property tax exemptions in the state of Texas. Also, the alternative property tax schedule from the multinomial logit estimation is defined by the property tax currently paid in the alternative houses selected from the choice set.

control function does not provide new relevant variation to the estimates. Given that house values in Texas are re-assessed every two to three years, moving costs represented by property taxes do not carry any information about historical house prices, and potential house buyers have similar user costs for the same property in that state.

Another interesting way of testing the fitness of the residential demand model is to look at predicted mobility rates. In my model, mobility rates can be derived from predicted probabilities of choosing a house, which are estimated directly from equation 8. Specifically, the estimated probabilities of choosing a house from the alternative set are used as a proxy for homeowners' mobility. Figure 11 plots these predicted probabilities by age groups. The predicted mobility patterns only resemble the ones showed in Figure 3 when using the model with the control function. When not using the variation in property taxes to control for unobserved components of the house, predicted mobility patterns are almost constant across age groups.

6. Conclusion

Given that homeownership is the primary way in which families accumulate wealth, understanding housing demand is of special importance in evaluating questions of welfare and equity across household types. Unfortunately, existing economic models that predict sorting in the urban landscape generally assume no barriers to household sorting due to transaction costs. In reality, it is hardly credible to assume that such frictions do not affect the housing market.

In this paper I provide clear evidence that transaction costs affect individual behavior. Using a natural experiment design generated by California's Propositions 13, 60 and 90, I show a distinct effect of property tax variations on household mobility. This analysis also indicates that individuals may face differentiated prices in the market, and that such variation can be used as a source of identification for revealed preference models of housing demand. MWTP for housing characteristics are approximately 100% upward biased when the variation in property taxes is not used to account for the endogeneity of housing prices. Given this evidence, economic models that fail to incorporate moving costs and other unobserved factors may provide biased predictions of choice behavior.

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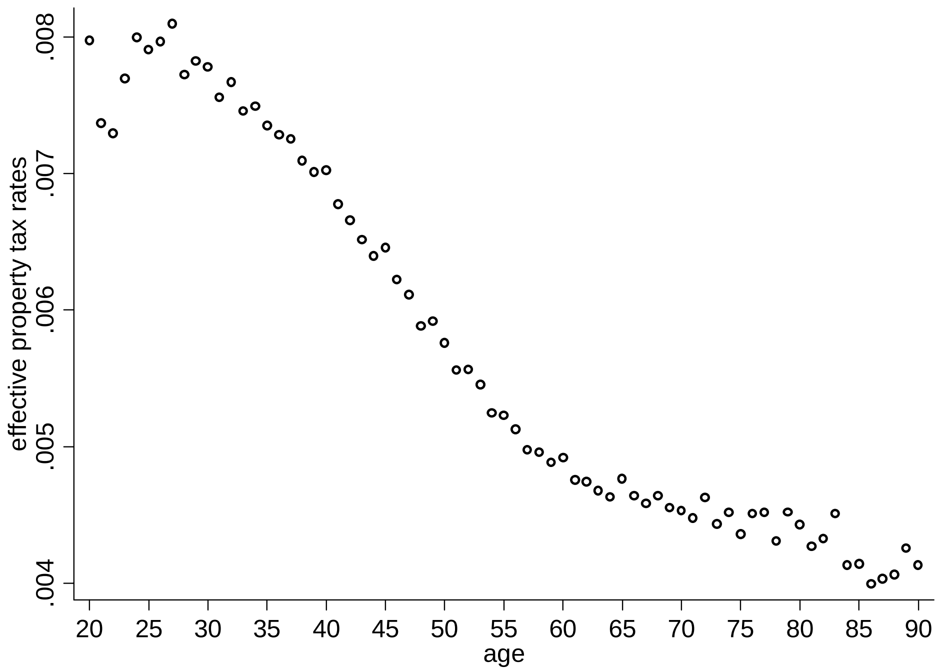
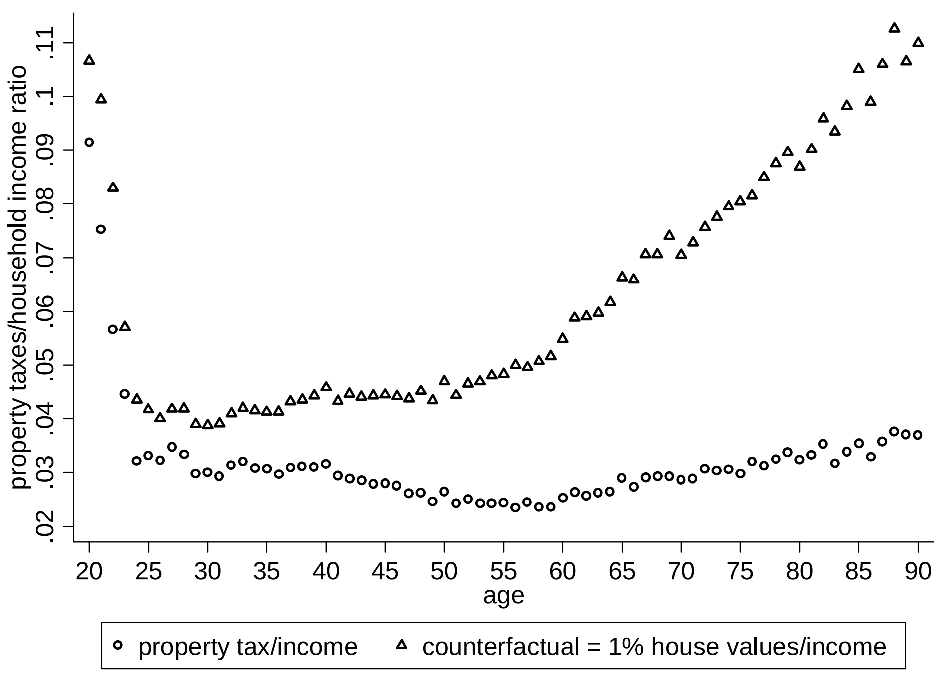
Figure 1. Effective property tax rates by homeowner age, California 1990.



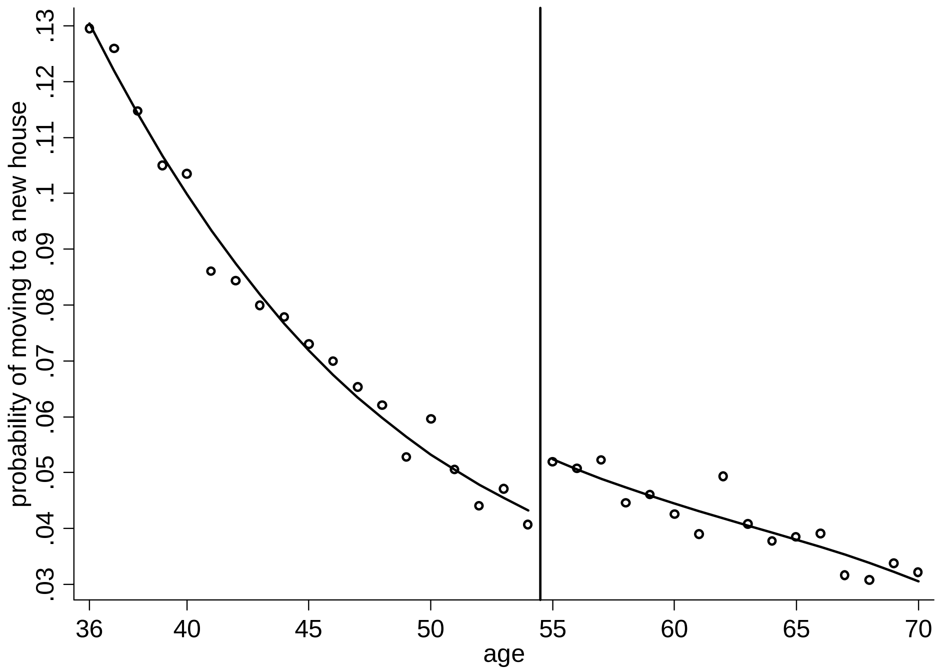
Source: 1990 IPUMS. Notes: Effective property tax rates are calculated as property taxes divided by house values. Age is the maximum between age of the head of the house and spouse.

Figure 2. Property taxes/household income ratio by age, California 1990.



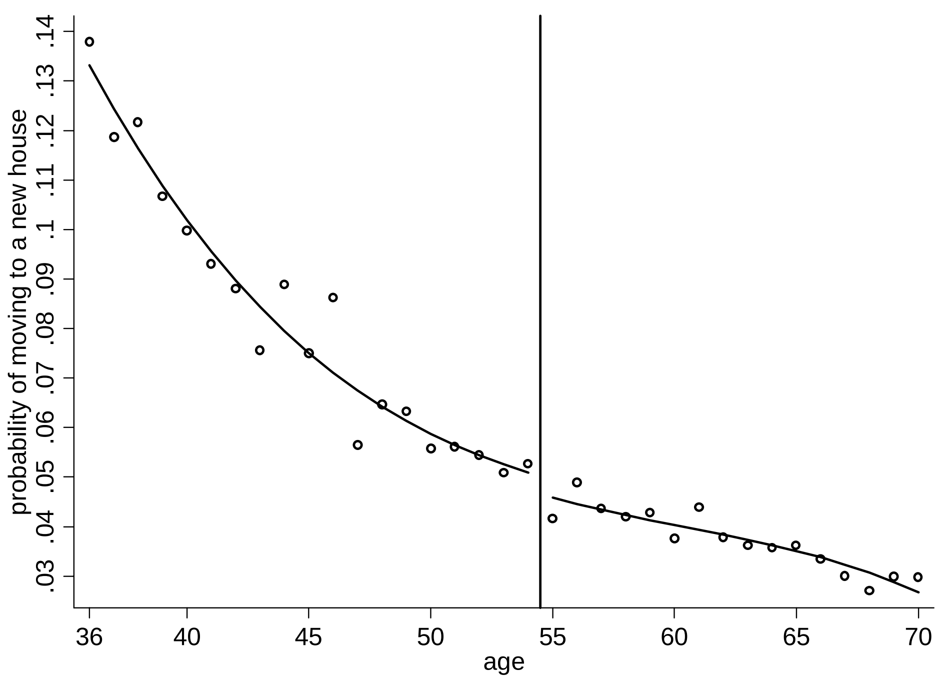
Source: 1990 IPUMS. Notes: Counterfactual ratios are calculated as 1% of house values divided by household income. Age is the maximum between age of the head of the house and spouse.

Figure 3. Probability of moving for homeowners by age, California 1990.



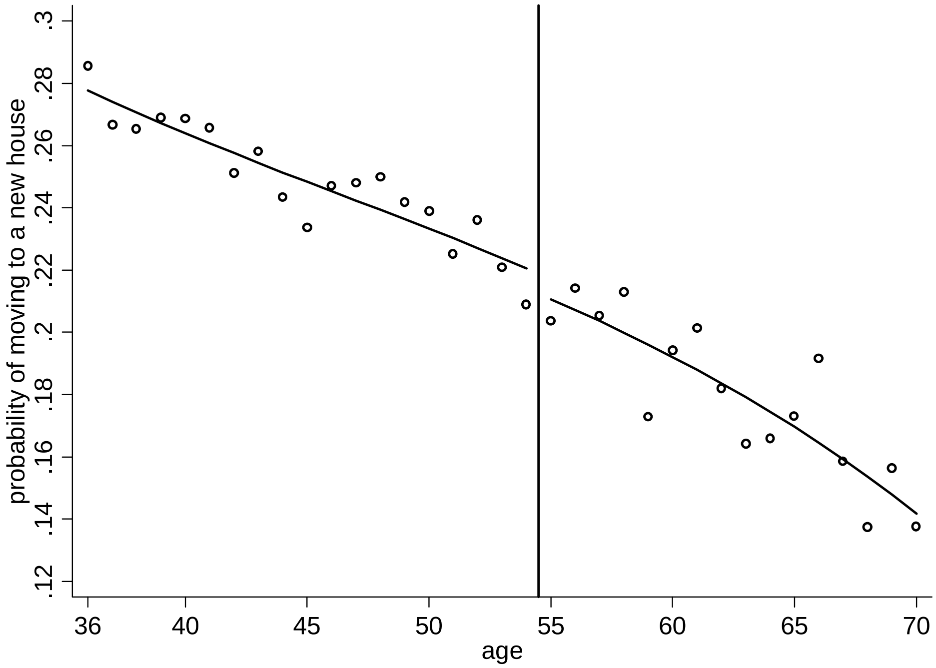
Source: 1990 IPUMS. Notes: Each dot represents the probability of moving for homeowners by age, calculated as the number of new movers in 1989-1990 divided by the total number of homeowners by age. Age is the maximum between age of the head of the house and spouse. The thick line is composed by predicted values of a polynomial regression of probability of moving on a dummy for 55-year, age, age squared and cubic, household characteristics and housing amenities.

Figure 4. Probability of moving for homeowners by age, California 1980.



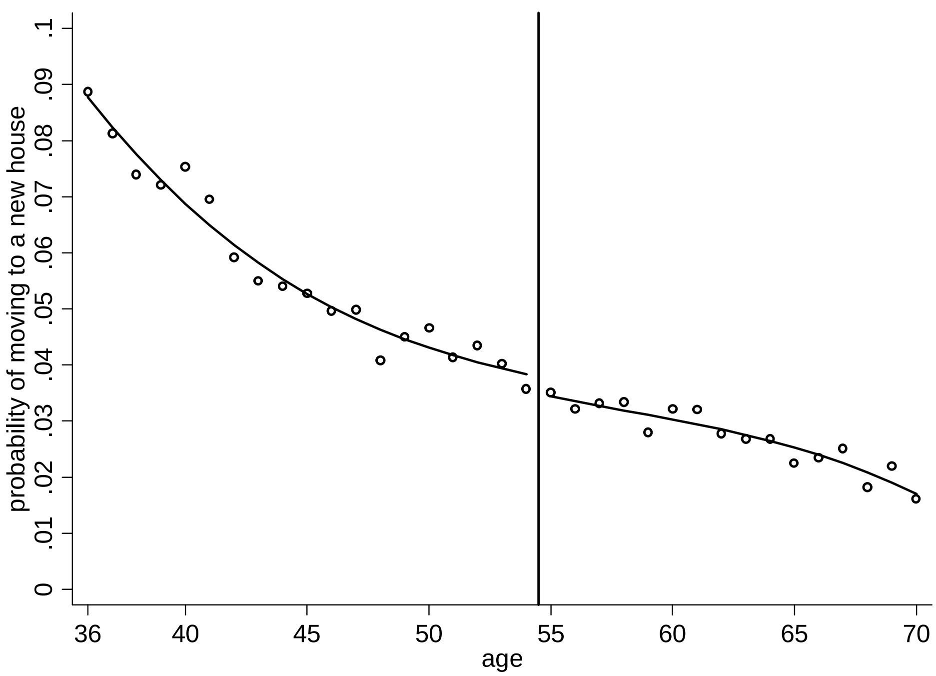
Source: 1980 IPUMS. Notes: Each dot represents the probability of moving for homeowners by age, calculated as the number of new movers in 1979-1980 divided by the total number of homeowners by age. See Figure 4 for other details.

Figure 5. Probability of moving for renters by age, California 1990.



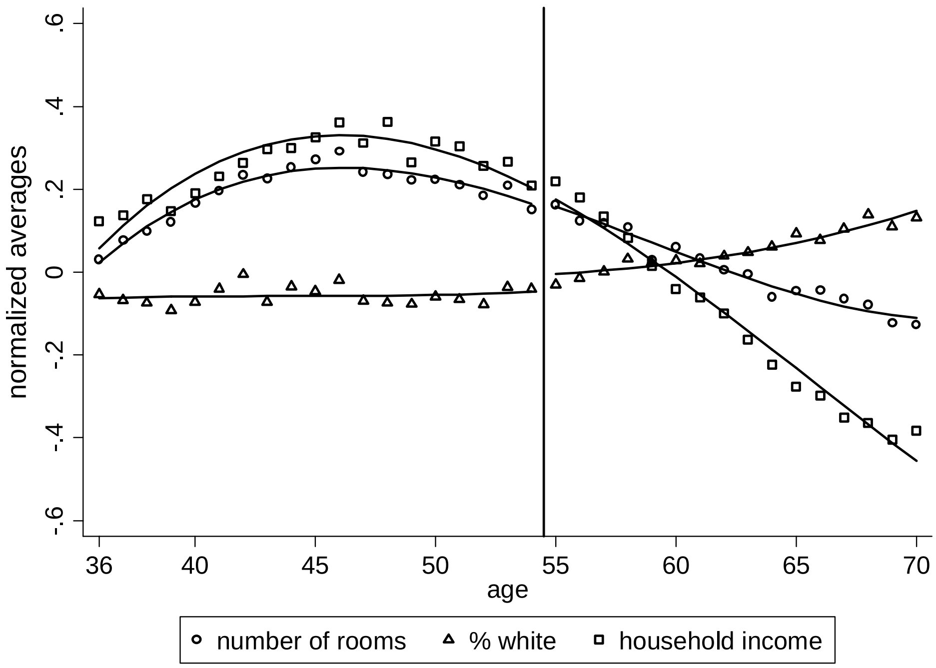
Source: 1990 IPUMS. Notes: Each dot represents the probability of moving for renters by age, calculated as the number of new renters in 1989-1990 divided by the total number of renters by age. See Figure 3 for details.

Figure 6. Probability of moving for homeowners by age, Texas 1990.



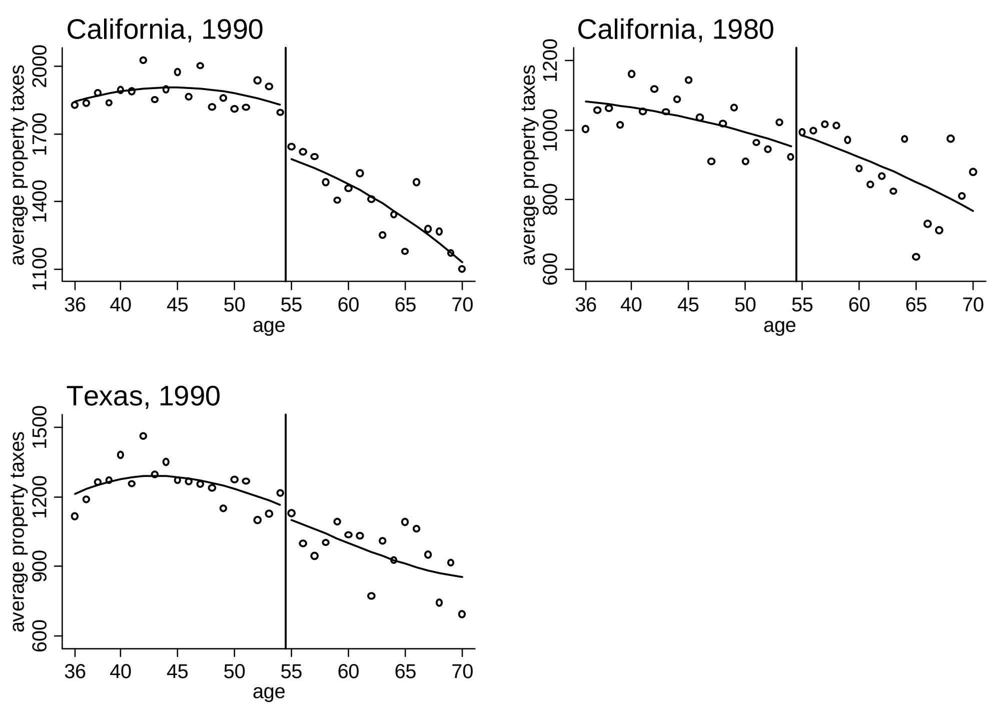
Source: 1990 IPUMS. See Figure 4 for details.

Figure 7. Other covariates by age, California 1990



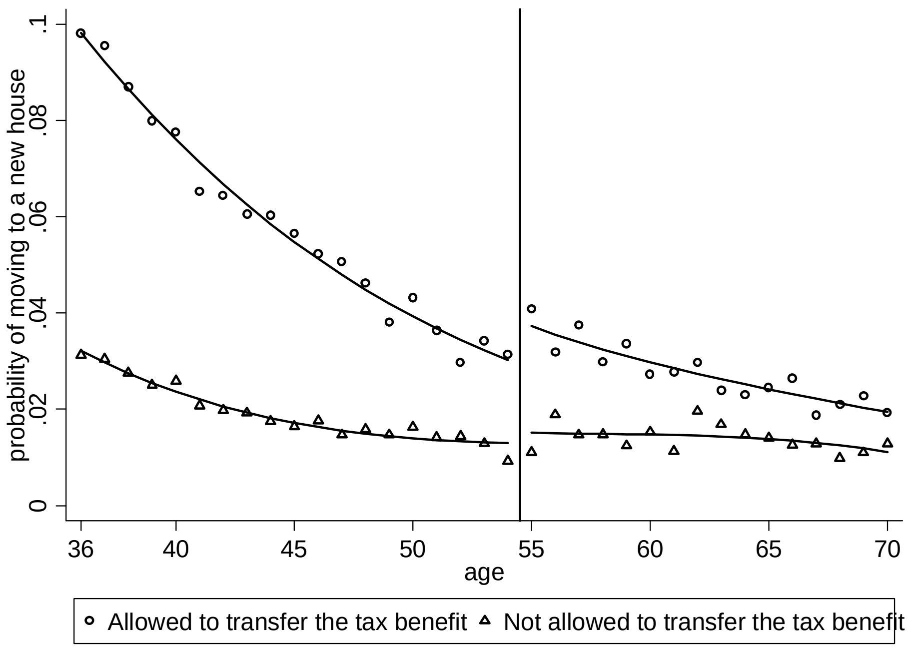
Source: 1990 IPUMS. Notes: Each dot represents the average normalized covariate by each age group. See Figure 4 for details.

Figure 8. Average property taxes for recent movers by age, California and Texas, 1990 and 1980.



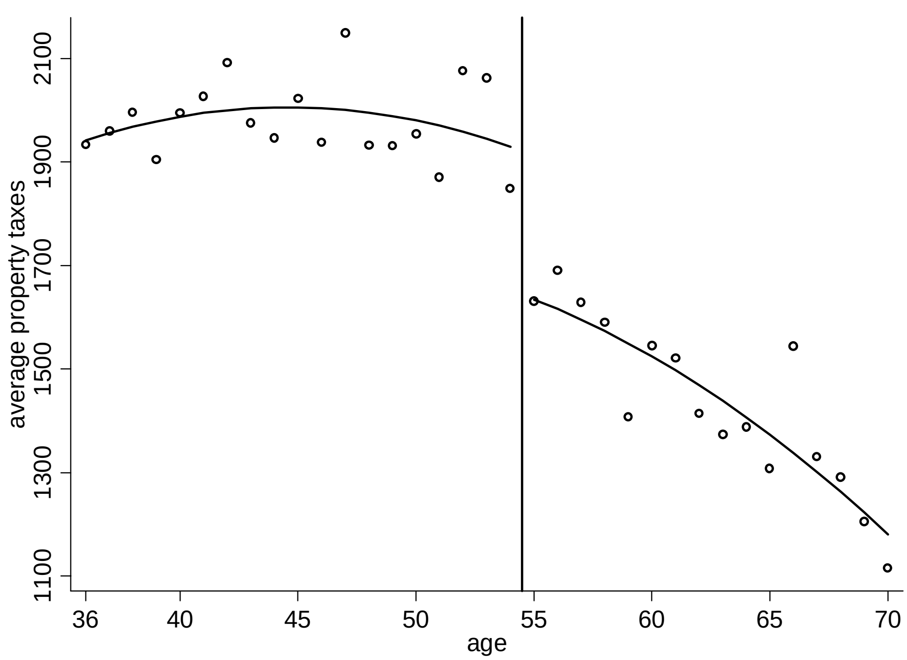
Source: 1990 IPUMS. Notes: Each dot represents the average property taxes faced by each age group. See Figure 4 for details.

Figure 9. Probability of moving for homeowners allowed and not allowed to transfer the tax benefit for homeowners by age, California 1990.



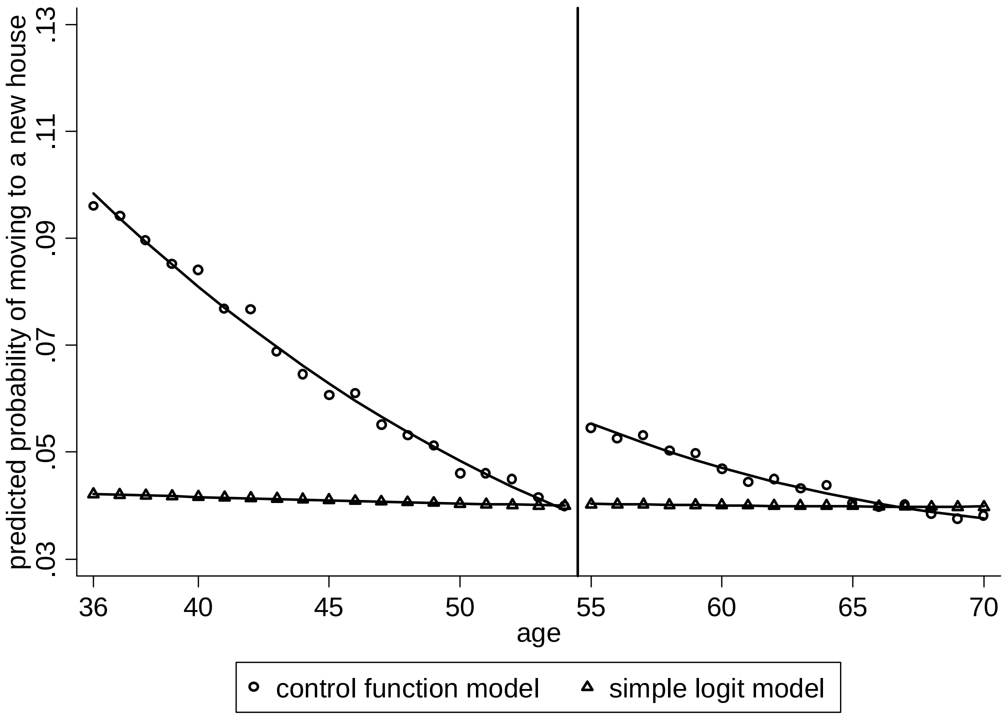
Notes: The group of homeowners allowed to transfer the tax benefit are the ones who moved in 1989-1990 within the same county, or to a different county that approved Proposition 90. See Figure 4 for other details.

Figure 10. Average property taxes for recent movers allowed to transfer the tax benefit by age, California 1990.



Notes: Each dot represents the average property taxes faced by each age group. See Figure 4 for details.

Figure 11. Predicted probabilities of moving by age, San Francisco Bay Area 1990



Notes: The predicted probabilities are normalized by the observed mobility rates of 54-year old homeowners, in order to make the predictions comparable to the results showed in Figure 4. The control function model includes all choice variables and heterogeneity.

Table 1. House values and property taxes by 'when moved in', California and Texas, 1990 and 1980.

	Full Sample	Moved in 1989	Moved in 1985-1988	Moved in 1980-1984	Moved in 1970-1979	Moved in 1960-1969	Moved in before 1960
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
California, 1990							
house value	215042	210820	219730	214524	219431	215399	194380
property tax	1179	1666	1672	1411	866	616	541
effective tax rate	**0.58%**	**0.80%**	**0.78%**	**0.70%**	**0.44%**	**0.34%**	**0.34%**
California, 1980							
house value	97151	105824	103473	101039	94977	83857	74627
property tax	774	931	950	738	681	586	517
effective tax rate	**0.85%**	**0.92%**	**0.96%**	**0.79%**	**0.78%**	**0.78%**	**0.80%**
Texas, 1990							
house value	74669	79589	83686	78008	72734	67323	56559
property tax	950	1080	1168	1024	928	757	475
effective tax rate	**1.39%**	**1.50%**	**1.49%**	**1.43%**	**1.40%**	**1.27%**	**1.06%**
Texas, 1980							
house value	51391	61642	60753	52337	46174	38442	30679
property tax	704	814	888	726	617	476	307
effective tax rate	**1.44%**	**1.39%**	**1.52%**	**1.48%**	**1.44%**	**1.34%**	**1.14%**

Source: 1990 and 1980 IPUMS. Notes: Household data include all 18-year or older head of the house homeowners, with non-zero property tax payments. Households with allocations for 'when moved in' and 'property taxes' are not included. Households living in another state 5 years prior to the relevant year of moving are also excluded. Half of the 1980 sample was not included because the Census Bureau did not process the mobility variables for a random sample of half of the population, in order to reduce costs of processing information. Number of observations is reported in appendix table 1.

Table 2. Reduced form estimates - effect of age on the probability of moving to another house in California 1990.

	California, 1990				pooling Texas, 1990	pooling California, 1980	not including 54 to 55-year olds
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1 if age>=55	**-0.0841**	**0.0547**	**0.0153**	**0.0152**	-0.0031	-0.0043	**0.0143**
	(0.0011)	(0.0092)	(0.0030)	(0.0029)	(0.0036)	(0.0047)	(0.0035)
age		-0.0051	-0.0241	-0.0230	-0.0223	-0.0287	-0.0209
		(0.0003)	(0.0015)	(0.0015)	(0.0016)	(0.0021)	(0.0040)
1 if California X 1 if age>=55					**0.0173**		
					(0.0050)		
1 if 1990 X 1 if age>=55						**0.0202**	
						(0.0061)	
age controls (squared and cubic)	N	N	Y	Y	Y	Y	Y
characteristics of the house	N	N	N	Y	Y	Y	Y
characteristics of the household	N	N	N	Y	Y	Y	Y
metropolitan area dummies	N	N	N	Y	N	N	Y
state fixed effects	N	N	N	N	Y	N	N
year fixed effects	N	N	N	N	N	Y	N
interactions of CA and age controls	N	N	N	N	Y	N	N
interactions of 1990 and age controls	N	N	N	N	N	Y	N
sample size	233514	233514	233514	233514	381069	308059	225411

Source: 1980 and 1990 IPUMS. Notes: Table shows maximum likelihood probit estimates (and standard errors) of the effect of age on the probability of moving to a new house in 1989-1990, evaluating the marginal effect at the mean. Age is the maximum between age of the head of the house and spouse. House characteristics include number of rooms and house value. Household characteristics include income, race, and education. Standard error estimates are based on the Eicker-White formula to correct for heteroskedasticity and they are also clustered by age.

Table 3. Effect of age on effective property tax rates in California 1990.

	Recent Movers in California, 1990				Long Tenure Homeowners CA, 1990	Recent Movers Texas, 1990	Recent Movers CA, 1980
	(1)	(2)	(3)	(4)	(5)	(7)	(8)
1 if age>=55	**-0.078**	**-0.087**	**-0.084**	**-0.083**	-0.008	0.0067	-0.003
	(0.008)	(0.020)	(0.020)	(0.020)	(0.005)	(0.0593)	(0.028)
age		-0.014	-0.014	-0.014	-0.023	-0.0025	0.013
		(0.008)	(0.008)	(0.008)	(0.004)	(0.0240)	(0.011)
age squared and cubic	N	Y	Y	Y	Y	Y	Y
characteristics of the house	N	N	Y	Y	Y	Y	Y
characteristics of the household	N	N	Y	Y	Y	Y	Y
metropolitan area dummies	N	N	N	Y	Y	Y	Y
R-squared	0.005	0.005	0.008	0.014	0.058	0.0535	0.018
sample size	19854	19854	19854	19854	112705	9118	7220

Source: 1980 and 1990 IPUMS. Notes: Table shows estimates (and standard errors) of a linear regression model of the effect of age on effective property tax rates for recent movers. Households who were living in another state 5 years prior to the relevant year of moving were excluded. Standard error estimates are based on the Eicker-White formula to correct for heteroskedasticity and they are also clustered by age. Estimates are presented in percentage points.

Table 4. Multinomial logit estimates, 30 to 70-year old homeowners, San Francisco Bay Area, 1990.

	simple logit	other house variables	other neigh. variables	all variables	control function	wealth effects	
						logit	control function
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
user cost	-0.4798	-0.4635	-0.5894	-0.7972	-2.4758	0.3348	13.1502
	(0.0055)	(0.0054)	(0.0055)	(0.0103)	(0.0170)	(0.0086)	(0.0653)
number of rooms	0.2974	0.3743	0.3272	0.5556	0.9790	0.3672	0.8782
	(0.0047)	(0.0046)	(0.0046)	(0.0083)	(0.0103)	(0.0077)	(0.0093)
1 if detached house	0.3749	0.1975	0.4284	0.3938	0.5070	0.2506	0.6877
	(0.0045)	(0.0046)	(0.0045)	(0.0102)	(0.0111)	(0.0101)	(0.0106)
average neighborhood income	0.1473	0.2038	0.1680	0.2755	0.8559	-0.0150	1.0531
	(0.0046)	(0.0045)	(0.0049)	(0.0083)	(0.0118)	(0.0074)	(0.0105)
control function					1.6928		12.7426
					(0.0074)		(0.0621)
house controls	N	Y	N	Y	Y	Y	Y
neighborhood controls	N	N	Y	Y	Y	Y	Y
number of households	98407	98407	98407	98407	98407	98407	98407

Source: US Census Bureau - 1990 California Decennial Census Long Form data. Notes: Table shows multinomial logit estimates (and standard errors) of the household location decision model. Other house variables include: built in 1985-1989, built in 1980-1984, and built in 1970-1979. Other neighborhood controls include: block group percentage white, density, elevation, air quality, and first grade test scores. Standard errors were adjusted by the following bootstrap procedure: a) repeatedly estimate the first stages with bootstrapped user cost samples, b) re-estimate the multinomial logit with the new residuals, c) calculate the sample variance of the estimated coeffcients and add it to the standard errors obtained from the traditional formulas.

Table 5. Implied marginal willingness to pay for housing characteristics, 30 to 70-year old homeowners, San Francisco Bay Area, 1990.

	simple logit	other house variables	other neigh. variables	all variables	control function	wealth effects	
						logit	control function
	(1)	(2)	(3)	(4)	(5)	(8)	(9)
1 extra room	3,275	4,327	2,975	3,994	2,021	34217	2083
	(65)	(75)	(51)	(79)	(32)	(1124)	(24)
detached house	17,210	9,325	15,903	9,224	4,446	96002	6706
	(290)	(248)	(227)	(268)	(117)	(4555)	(108)
$10,000 change in neighborhood income	1,081	1,535	995	1,236	1,298	-905	1616
	(36)	(39)	(31)	(41)	(27)	(446)	(18)
house controls	N	Y	N	Y	Y	Y	Y
neighborhood controls	N	N	Y	Y	Y	Y	Y
average user cost	20035	20035	20035	20035	20035	20035	20035
number of households	98407	98407	98407	98407	98407	98407	98407

Source: US Census Bureau - 1990 California Decennial Census Long Form data. Notes: Table shows MWTP estimates (and standard errors) derived from Table 4. Standard errors were calculated using the delta method. MWTP estimates for the wealth effects were adjusted to the average user cost of the house.

Table 6. Implied marginal willingness to pay for housing characteristics - heterogeneity in preferences.

	Baseline	MWTP Income (+$10,000)	MWTP Age (+10 years)	MWTP White (vs other races)
	(1)	(2)	(3)	(4)
1 extra room	1,783	1,832	1,801	1,826
	(49)	(53)	(52)	(54)
detached house	3,509	3,468	2,339	3,243
	(140)	(141)	(95)	(138)
$10,000 change in neigh income	1,532	1,574	1,426	1,523
	(36)	(38)	(35)	(37)
10% increase in %white	256	259	218	469
	(28)	(29)	(26)	(49)
house controls	Y	Y	Y	Y
neighborhood controls	Y	Y	Y	Y
average user cost	20035	20035	20035	20035

Source: US Census Bureau - 1990 California Decennial Census Long Form data. Notes: Table shows MWTP estimates (and standard errors) derived from a multionomial logit model that includes heterogeneith, ie., all choice variables were interected with individual income, a dummy for white, age and distance to work. The baseline estimates correspond to a mixed race household with average income and average age. Columns (2) to (4) report willingness to pay associated with a household listed in the column heading, holding all other factors equal. Other house variables include: built in 1985-1989, built in 1980-1984, and built in 1970-1979. Other neighborhood controls include: block group percentage white, density, elevation, air quality, and first grade test scores. Standard errors were calculated using the delta method.

Table 7. Multinomial logit estimates, 30 to 64-year old homeowners, Dallas, TX 1990.

	Mulinomial Logit Estimates simple logit	Mulinomial Logit Estimates control function	MWTP simple logit	MWTP control function
	(1)	(2)	(3)	(4)
user cost	-0.1581	-0.1860		
	(0.0164)	(0.0223)		
number of rooms	0.0831	0.0989	2,164	2,188
	(0.0147)	(0.0173)	(444)	(464)
1 if detached house	0.0392	0.0397	6,404	5,516
	(0.0115)	(0.0116)	(2000)	(1741)
average neighborhood income	-0.0669	-0.0600	-2,863	-2,183
	(0.0118)	(0.0124)	(585)	(520)
control function		0.0314		
		(0.0272)		
house controls	Y	Y	Y	Y
neighborhood controls	Y	Y	Y	Y
average user cost			8242	8242
number of households	9970	9970	9970	9970

Source: 1990 IPUMS. Notes: Table shows multinomial logit estimates (and standard errors) of the household location decision model. Other house variables built in 1985-1989, built in 1980-1984, and built in 1970-1979. Other neighborhood controls include: neighborhood percentage white. Standard errors for the function approach are adjusted by the following bootstrap procedure: a) estimate 1000 first stages with bootstrapped user cost samples, b) re-estimate the mu logit with the new residuals, c) calculate the sample variance of the estimated coeffcients and add it to the standard errors obtained from the traditional form

Appendix Table 1. Summary of Census data by 'when moved in', California and Texas, 1990 and 1980.

	Full Sample (1)	Moved in 1989 (2)	Moved in 1985-1988 (3)	Moved in 1980-1984 (4)	Moved in 1970-1979 (5)	Moved in 1960-1969 (6)	Moved in before 1960 (7)
California, 1990							
number of rooms	5.99	5.77	5.93	5.99	6.17	6.12	5.71
1 if detached	86.2%	76.6%	80.2%	83.1%	88.9%	95.9%	96.4%
1 if built in 1960-70	21.0%	18.3%	20.3%	24.6%	39.4%	0.0%	0.0%
1 if built before 1960	58.8%	33.6%	37.9%	46.6%	60.6%	100.0%	100.0%
age	52	41	43	48	54	63	71
household income	58183	62840	64338	62885	59146	50879	36671
1 if white	88.3%	86.1%	87.0%	87.7%	88.9%	89.6%	91.7%
1 if black	3.8%	2.4%	2.9%	3.1%	4.5%	5.5%	5.0%
1 if college or more	64.7%	73.1%	74.0%	70.8%	64.0%	52.5%	39.1%
sample size	233514	19854	66581	34374	59310	30092	23303
California, 1980							
number of rooms	6.04	6.13	6.16	6.17	6.07	5.64	5.50
1 if detached	97.5%	97.5%	97.5%	97.6%	97.9%	97.5%	95.1%
1 if built in 1960-70	23.3%	18.0%	21.5%	25.9%	43.9%	0.0%	0.0%
1 if built before 1960	54.8%	32.9%	37.5%	47.3%	56.1%	100.0%	100.0%
age	50	39	41	47	55	62	71
household income	27998	29964	29965	29504	28433	23443	17213
1 if white	90.0%	89.9%	89.3%	89.1%	90.0%	91.6%	94.5%
1 if black	4.7%	3.2%	3.9%	5.7%	5.5%	4.8%	3.6%
1 if college or more	50.2%	62.6%	61.2%	53.2%	43.9%	33.4%	25.9%
sample size	74545	7220	21823	14210	17535	9986	3771
Texas, 1990							
number of rooms	5.94	5.91	5.98	5.90	6.06	5.95	5.63
1 if detached	90.1%	86.4%	85.4%	85.1%	92.4%	97.9%	98.3%
1 if built in 1960-70	25.1%	20.4%	19.2%	24.1%	54.8%	0.0%	0.0%
1 if built before 1960	47.7%	29.2%	23.9%	27.8%	45.2%	100.0%	100.0%
age	52	40	42	47	54	63	71
household income	43637	46944	49166	48002	43863	37134	26641
1 if white	93.1%	93.4%	93.6%	94.3%	93.0%	91.8%	91.4%
1 if black	5.5%	4.6%	4.5%	4.0%	6.0%	7.6%	8.3%
1 if college or more	52.4%	60.8%	63.4%	59.1%	51.0%	38.1%	27.5%
sample size	147555	9118	37338	29713	38082	18220	15084
Texas, 1980							
number of rooms	6.00	6.04	6.19	6.11	5.96	5.61	5.47
1 if detached	98.8%	98.5%	98.2%	98.9%	99.4%	99.2%	99.2%
1 if built in 1960-70	24.7%	15.1%	18.1%	26.7%	55.5%	0.0%	0.0%
1 if built before 1960	43.7%	23.0%	23.0%	32.1%	44.5%	100.0%	100.0%
age	49	38	40	48	55	62	71
household income	25373	27863	28131	26585	24903	21122	13681
1 if white	92.6%	95.0%	93.7%	92.3%	91.6%	91.7%	89.2%
1 if black	6.2%	3.5%	4.8%	6.6%	7.3%	7.4%	9.9%
1 if college or more	44.2%	59.2%	57.5%	44.8%	35.3%	26.5%	21.1%
sample size	38146	3307	11938	7243	8686	4793	2179

Source: 1980 and 1990 IPUMS. Notes: Household data include all 18-year or older head of the house homeowners, with non-zero property tax payments. Households with allocations for 'when moved in' and 'property taxes' are not included. Households living in another state 5 years prior to the relevant year of moving are also excluded. Half of the 1980 sample was not included because the Census Bureau did not process the mobility variables for a random sample of half of the population, in order to reduce costs of processing information.

Appendix Table 2. Multinomial logit sample, San Francisco Bay Area 1990.

	30 to 70-year old homeowners (1)	choice set (2)
house value	306984	301609
number of rooms	6.31	5.92
1 if detached	0.85	0.72
1 if built in 1985-89	0.08	0.27
1 if built in 1980-84	0.07	0.09
1 if built in 1970-79	0.20	0.20
block group average income	64546	63941
block group % white	0.73	0.72
elevation	248.00	250.00
1st grade test scores	545.00	541.00
population density	0.28	0.27
air quality index	23.29	23.47
sample size	98407	8347

Source: US Census Bureau - 1990 California Decennial Census Long Form data and 1990 IPUMS. Notes: Average income, percentage white, and density are constructed at the block group level. Elevation is measured at the block level (source: EPA: BASINS). Test scores are assigned from the closest school within the school district (source: California Department of Education, 1991-1993). Air quality is predicted for each census block using information from monitor stations (source: Rand California, 1990) and industrial plants (source: EPA – AIRS –Aerometric Information Retrieval System, 1990).

Appendix Table 3. First stage estimates

	San Francisco, 30-70 year olds			Dallas, 30 to 64-year olds
	control function		wealth effect	control function
dependent variable: user cost	(1)	(2)	(3)	(4)
property taxes	4.2	4.0	-0.0003	4.0
	(0.004)	(0.004)	(0.000003)	(0.010)
number of rooms	801	792	-0.022	875
	(3.5)	(3.5)	(0.003)	(7.8)
1 if detached house	1,823	1,916	-0.146	-210
	(13.1)	(14.2)	(0.011)	(40)
average income of neighborhood	680	599	-0.032	574
	(2.3)	(2.6)	(0.002)	(11.2)
other house controls	N	Y	Y	Y
other neighborhood controls	N	Y	Y	Y
R-squared	0.83	0.84	0.04	0.77

Source: US Census Bureau - 1990 California Decennial Census Long Form data. Notes: Table shows multiple regression estimates of the user cost of the house on the instrumental variable and other housing and neighborhood characteristics. Other house variables include: built in 1985-1989, built in 1980-1984, and built in 1970-1979. Other neighborhood controls include: block group percentage white, density, elevation, air quality, and first grade test scores.

Appendix Table 4. Hedonic price regressions, 30 to 70-year old homeowners, San Francisco Bay Area, 1990.

	dependent variable: user cost			
	(1)	(2)	(3)	(4)
1 extra room	1,693	1,698	1,570	1,549
	(14)	(14)	(13)	(13)
detached house	1,149	1,328	1,762	2,434
	(59)	(62)	(57)	(60)
$10,000 change in neighborhood income	1,635	1,652	1,346	1,327
	(8)	(8)	(9)	(9)
house controls	N	Y	N	Y
neighborhood controls	N	N	Y	Y
R-squared	0.51	0.52	0.56	0.57
average user cost	20035	20035	20035	20035

Source: US Census Bureau - 1990 California Decennial Census Long Form data. Notes: Table shows linear regression estimates of the user cost of the house on housing and neighborhood characteristics. Other house variables include: built in 1985-1989, built in 1980-1984, and built in 1970-1979. Other neighborhood controls include: block group percentage white, density, elevation, air quality, and first grade test scores.