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You can take it with you: Proposition 13 tax benefits, residential mobility, and willingness to pay for housing amenities☆

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ABSTRACT

The endogeneity of prices has long been recognized as the main identification problem in the estimation of marginal willingness to pay (MWTP) for the characteristics of a given product. This issue is particularly important in the housing market, since a number of housing and neighborhood features are unobserved by the econometrician. This paper proposes the use of a well defined type of transaction costs–moving costs generated by property tax laws–to deal with this type of omitted variable bias. California's Proposition 13 property tax law is the source of variation in transaction costs used in the empirical analysis. 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. Its importance to homeowners is estimated from a natural experiment created by two amendments that allow households headed by an individual over the age of 55 to transfer the implicit tax benefit to a new home. Indeed, 55-year old homeowners have 25% higher moving rates than those of comparable 54 year olds. These transaction costs from the property tax laws are then incorporated into a household sorting model. The key insight is that because of the property tax laws, different potential buyers may have different user costs for the same house. The exogenous property tax component of this user cost is then used as an instrumental variable. I find that MWTP estimates for housing characteristics are approximately 100% upward biased when the choice model does not account for the price endogeneity.
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1. Introduction

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 that 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 cost–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. Estimation of marginal willingness to pay (MWTP) for housing and neighborhood amenities have been plagued by omitted variable bias, since finding a credible research design that accounts for the unobservable characteristics of a house is a daunting task.[4] The key insight of the instrumental variable developed in this paper 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 of about 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 revenues.[5]

☆ The first version of this paper circulated in 2004 under the title "You Can Take It with You: Transferability of Proposition 13 Tax Benefits, Residential Mobility, and Willingness to Pay for Housing Amenities".

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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 et al. (1984, 1993), Benabou (1993), Fernandez and Rogerson (1996), Nechyba (1997) and Epple and Sieg (1999).

[2] Recent examples are found in Barrow and Rouse (2004), Rothstein (2006) and Bajari and Kahn (2005).

[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 (2002) for a survey about the different types of housing transaction costs.

[4] Credible instruments to control for the correlation between prices and unobserved housing quality are scarce in this literature. Bajari and Kahn (2005) , for example, estimate bounds on willingness to pay for distance to work in order to avoid the use of instruments.

[5] 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), Silva and Sonstelie (1995), Fischel (1989), and Brunner and Rueben (2001) for fiscal consequences of Proposition 13.

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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 house may experience a large increase in tax liability.[6]

The importance of Proposition 13 to household mobility is estimated using a pair of amendments that passed in the late 1980's — Proposition 60 in 1986, and Proposition 90 in 1988. Those propositions allowed homeowners aged 55 or older who sell a property and buy another of equal or lesser value, 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. The lock-in effect attributable to Proposition 13 is estimated by comparing householders who are 54 years old to those who are 55, in a typical Regression Discontinuity (RD) design. 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 and its amendments are then explicitly incorporated in a household 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 et al. (1995), and Petrin and Train (2006).

The household sorting model developed in this paper has two distinct features. 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. Homeowners living in a house for long periods of time have much lower property taxes than recent movers, and, moreover, homeowners face different user costs when deciding to move due to Propositions 60 and 90. These institutional features of Propositions 13, 60 and 90 are then used to form an instrumental variable to deal with the endogeneity of housing prices. The intuition of the instrumental variable is that the property tax component of the user cost of the house is not endogenously determined as the housing price component, and therefore is uncorrelated with characteristics of the house or neighborhood.[7]

Another distinct characteristic of this study is that all estimates are fully based on micro-data. In fact, 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 workplace 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 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. The instrumental variable based on Proposition 13 and its amendments is used to solve this problem. Preference parameter estimates from this adjusted model are then used to recover MWTP for housing and neighborhood attributes. I find that most MWTP estimates for housing and neighborhood characteristics are approximately 100% upward biased when not controlling for the endogeneity of housing prices. For example, homeowners are willing to pay, on average, $1562 for a house in a neighborhood that has a marginal increase in the proportion of whites in a block group, compared to $2985 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 — MWTP estimates from models with or without the control function have identical results.

The large magnitude of differences in MWTP before and after dealing with the endogeneity of housing prices indicate that such problem is a first order concern in empirical applications, even in models that impose more structure to the estimation[8]. Although most characteristics of housing markets are arguably endogenous, complicating the search for a credible research design, estimates presented here indicate that property tax rules and other transaction costs can be used to credibly recover important structural parameters.[9]

The rest of this paper is organized as follows: In Section 2, I explain Proposition 13 in detail, and in Section 3, I estimate the importance of Proposition 13 to household mobility. Section 4 presents a household residential location model, and Section 5 provides estimates of MWTP for housing and neighborhood characteristics. Section 6 concludes.

2. 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 (1989) argues that the cost of the equalization program provoked a reaction by the voters in the form of restricting government revenues through Proposition 13.[10]

[6] These calculations are explained in Section 2. The Proposition 13 effect is analogous to the spatial lock-in related to falling housing prices, as in Caplin et al. (1997) and Chan (2001), or due to increase in interest rates, as in Quigley (1987).

[7] Section 4.2 explains in detail the parametric transformation of the dynamic features of these propositions into an instrumental variable.

[8] This issue has long been a concern in hedonic models — see Chay and Greenstone (2005) for a review of this problem in the WTP for air quality case.

[9] The direct impact of those property tax costs have also gained attention in the literature, such as in Farnham and Sevak (2006) which tests the Tiebout model using cross-state tax rules and empty nesters, and Shan (2008) which studies the direct impact of property tax rules on elderly mobility.

[10] Proposition 13' effects still reverberate today, as recurrent state budget pressures lead to underfunding of several public services. Paul Krugman wrote in a 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 $2264 on his $4 million home in Orange County."

Table 1
House values and property taxes by length of residence, California and Texas.

	Full sample	Moved in 1 year ago	Moved in 2–5 years ago	Moved in 6–10 years ago	Moved in 11–20 years ago	Moved in 21–30 years ago	Moved in 30 or more years ago
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
California, 1990							
House value	215,042	210,820	219,730	214,524	219,431	215,399	194,380
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	97,151	105,824	103,473	101,039	94,977	83,857	74,627
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	74,669	79,589	83,686	78,008	72,734	67,323	56,559
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	51,391	61,642	60,753	52,337	46,174	38,442	30,679
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.

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% 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.[11]

Table 1 compares the impact of Proposition 13 on effective tax rates (property taxes divided by house values) by the date when households moved into their home using the Integrated Public Use Microdata Series (IPUMS) 5% samples from 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 2 to 3 years in that state, and Texas is also a big state, with a large share of immigrants, and also from the Sunbelt. 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 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.

Fig. 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 (Fig. 2) the distributional effects of Proposition 13 are less pronounced. The main characteristic is that individuals between 50 and 60 years of age pay less property 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.

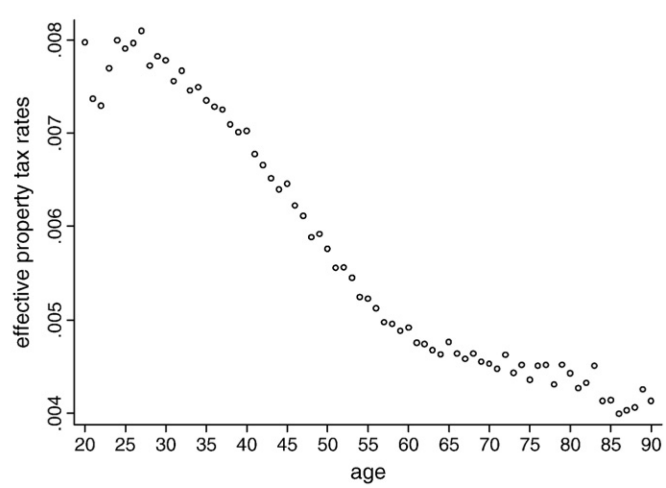

Fig. 1. Effective property tax rates by homeowner age, California 1990.
Source: 1990 IPUMS. Notes: Effective property tax rates are calculated as property taxes divided by house values. Age is the maximum between age of the head of the house and spouse.

[11] 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.

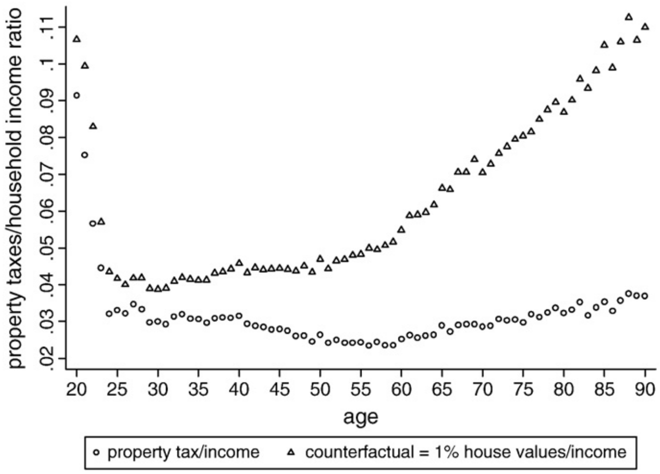

Fig. 2. Property taxes/household income ratio by age, California 1990.
Source: 1990 IPUMS. Notes: Counterfactual ratios are calculated as 1% of house values divided by household income. Age is the maximum between age of the head of the house and spouse.

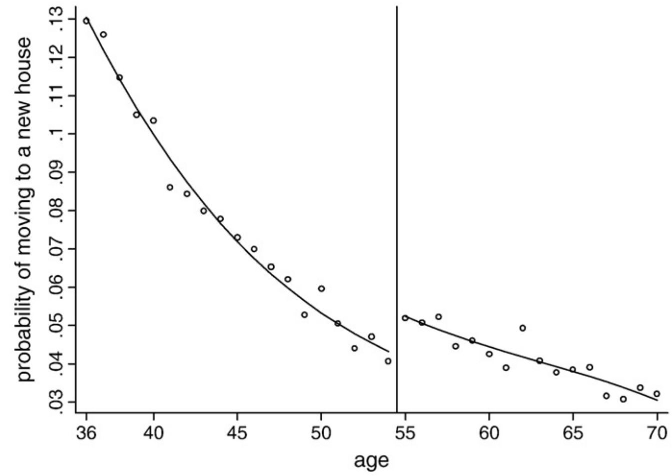

Fig. 3. Probability of moving for homeowners by age, California 1990.
Source: 1990 IPUMS. Notes: Each dot represents the probability of moving for homeowners by age, calculated as the number of new movers in 1989–1990 divided by the total number of homeowners by age. Age is the maximum between age of 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.

3. Lock-in effect

Do households in California respond to the property tax variation generated by Proposition 13? We are interested in the answer to this question for two reasons. First, it will indicate whether or not this type of transaction cost has enough relevance to be included in a residential location decisions model. Second, property taxes can only be used as an instrumental variable to solve the price endogeneity problem if both the tax benefit due to Proposition 13 is large and homeowners indeed respond to that variation in property taxes.

One way to estimate the lock-in effect from Proposition 13 is by comparing mobility rates of California to other states that have different tax laws. However, comparisons across states may be problematic, since different property tax systems can be correlated with other underlying characteristics of the state that also affect household mobility.[12] To circumvent this issue I make use of two important modifications to the California property tax law that were enacted during the 1980's — Proposition 60 and Proposition 90.

Proposition 60 was a constitutional amendment approved in 1986, which allowed the transfer of tax benefits for within-county movers. It 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 olds or older households to carry the frozen property tax payment to a new home within the same county. Proposition 90, approved in 1988, resulted in even more flexibility, allowing inter-county base year value transfers. Adoption of Proposition 90 was not mandatory and the law only applies across counties that approved the ordinance.[13]

I identify the impact of Proposition 13 on mobility by examining moving rates after the age 55, when the lock-in effect is removed. Fig. 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 Fig. 3 is calculated as the total number of homeowners that moved in the last year, divided by the total number of homeowners from the respective age.[14] 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 figure presented above is typical of a Regression Discontinuity (RD) approach. Although the RD is not new–see Thistlethwaite and Campbell (1960) and Cook and Campbell (1979)–its different estimation approaches were only developed recently. Hahn et al. (2001), for example, focused on identification and nonparametric estimation of the RD when the discontinuity occurs on a continuous variable. Lee (2001, 2008) and Lee and Card (2008) show that when the variable on which the discontinuity occurs is discrete, the RD can be estimated parametrically. Because age is measured in years we use a parametric rather than a nonparametric approach. The following third order polynomial equation for the probability of moving in 1990 is estimated:

$$Pr(\text{moving})_i = \Phi\left(\delta_1 D_i^{55} + \delta_2 \text{Age}_i + \delta_3 \text{Age}_i^2 + \delta_4 \text{Age}_i^3\right) \quad (1)$$

where D_i^{55} is a dummy for 55 years 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.[15]

[12] Wasi and White (2005) use such strategy. Other papers that study the effects of Proposition 13 on mobility are Sexton et al. (1995) and (1999). Nagy (1997) looks at California mobility rates before and after the law approval, finding no significant effects. This is understandable since the lock-in effect could only have an impact after a significant house prices increase.

[13] 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. 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.

[14] The fitted line is generated from a set of predicted values from a polynomial regression explained below — see Eq. (1).

[15] Interactions of age with post age 55 are not included because there are not enough degrees of freedom by year.

Table 2
RDD estimates of the effect of Propositions 60 and 90 on mobility.

	California, 1990			Pooling TX, 1990	Pooling CA, 1980	Excluding 54 to 55-year olds
Dependent variable: 1 if moved	(1)	(2)	(3)	(4)	(5)	(6)
1 if age ≥ 55	−0.084	0.015	0.015	−0.003	−0.004	0.014
	−0.001	−0.003	−0.003	−0.004	−0.005	−0.003
Age		−0.024	−0.023	−0.022	−0.029	−0.021
		−0.002	−0.002	−0.002	−0.002	−0.004
1 if California X 1 if age ≥ 55				0.017		
				−0.005		
1 if 1990 X 1 if age ≥ 55					0.020	
					−0.006	
Age controls (squared and cubic)	N	Y	Y	Y	Y	Y
Housing and household controls	N	N	Y	Y	Y	Y
Metropolitan area dummies	N	N	Y	N	N	Y
State fixed effects	N	N	N	Y	N	N
Year fixed effects	N	N	N	N	Y	N
Sample size	233,514	233,514	233,514	381,069	308,059	225,411

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 2 reports results from a probit model, designed to quantify the patterns observed in the figures above. 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 (2) 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 RD: Covariates should not matter around the threshold that defines treatment and control groups.

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-year old 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.[16]

These reduced form results hold for the local population of 54 and 55-year olds, independent of their moving status. Given the 1-year difference in both cohorts, there is no reason to expect differences in preferences or average characteristics of those households. Fig. 4 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 to a causal relationship between the ability to transfer the tax benefit and mobility rates. The next subsection presents a variety of tests of this interpretation.

3.1. Robustness checks

In order to rule out competing hypotheses, I compare 1990 California data with several control groups, such as California data from 1980, before Propositions 60 and 90 had been approved. Fig. 5 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.

Fig. 6 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. Fig. 7 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 year olds in California in 1990.[17]

The Proposition 60 and Proposition 90 mechanism 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. We test it in Fig. 8 by comparing average property taxes by age for three different samples. The gap between 54 and 55 year old recent movers 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 estimated separately for recent movers and long tenure homeowners. Those differences are only significant for recent movers in California, 1990.[18]

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 would be $536 per

[16] 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). On the other hand, mobility rates from the Census 2000 show that a much smaller effect resulted from those amendments.

[17] 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 are 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).

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

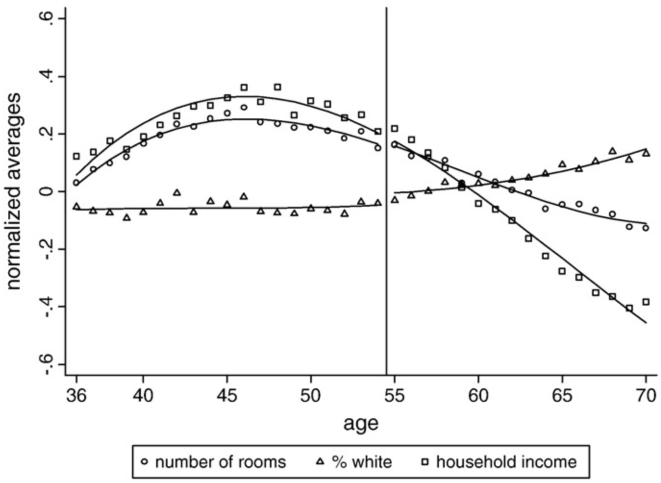

Fig. 4. Other covariates by age, California 1990.
Source: 1990 IPUMS. Notes: Each dot represents the average normalized covariate by each age group. Each thick line is composed by predicted values of a polynomial regression of the normalized covariate on a dummy for 55-year, age, age squared and cubic.

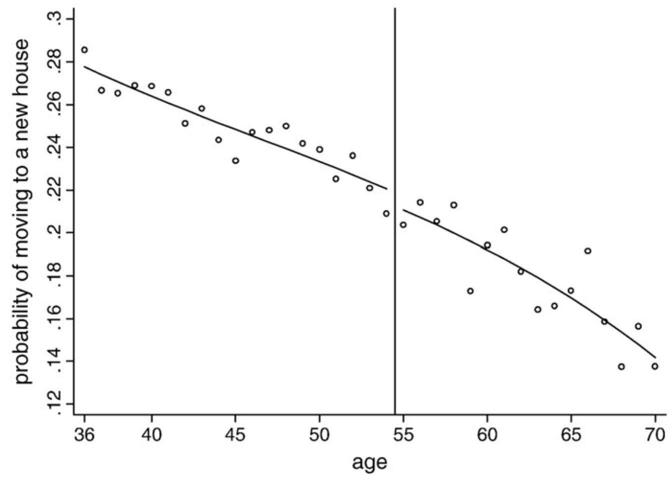

Fig. 6. Probability of moving for renters by age, California 1990.
Source: 1990 IPUMS. Notes: Each dot represents the probability of moving for renters by age, calculated as the number of new renters in 1989–1990 divided by the total number of renters by age. The thick line is composed by predicted values of a polynomial regression of probability of moving on a dummy for 55-year, age, age squared and cubic, household characteristics and housing amenities.

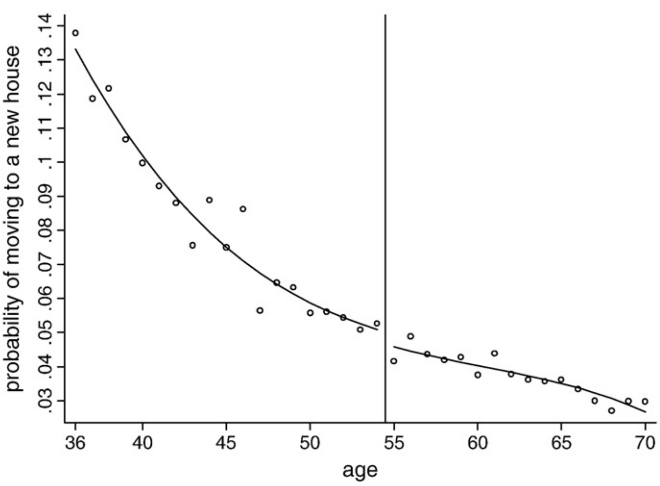

Fig. 5. Probability of moving for homeowners by age, California 1980.
Source: 1980 IPUMS. Notes: Each dot represents the probability of moving for homeowners by age, calculated as the number of new movers in 1979–1980 divided by the total number of homeowners by age. The thick line is composed by predicted 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.

year.[19] 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. Fig. 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 old recent homeowners moved to places that did not accept Proposition 90. The figure shows a close to zero impact on mobility for homeowners who could not transfer the tax benefit.[20]

4. Residential location decisions model

In this section I develop a household residential demand model, where the taxation costs explained in previous sections are included in the model as a device to recover estimates of the MWTP for housing and neighborhood attributes. The model is based on standard differentiated product demand models, whose roots lie in the work of McFadden (1974; 1978) and more recently Berry (1994), Berry et al. (1995), and Petrin and Train (2006). 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. The details of the model are described in Section 4.1, while Section 4.2 explains the endogeneity of housing prices and the research design related to property taxation that deals with this problem.

4.1. 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 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

[19] In order to estimate the full tax benefit, one would 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.

[20] 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.

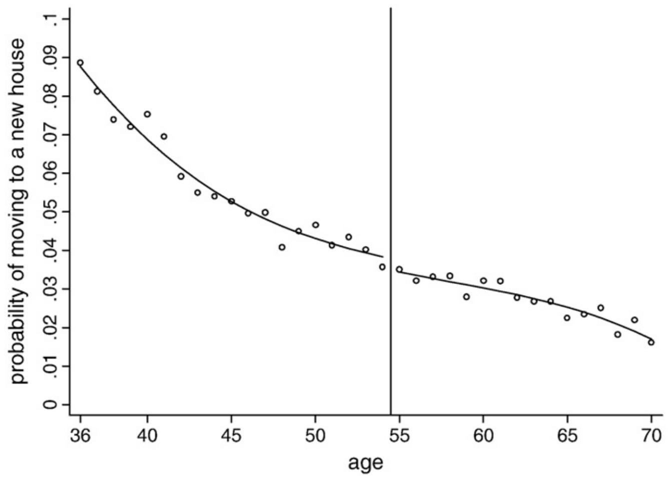

Fig. 7. Probability of moving for homeowners by age, Texas 1990.
Source: 1990 IPUMS. Notes: Each dot represents the probability of moving for homeowners by age, calculated as the number of new movers in 1989–1990 divided by the total number of homeowners by age. The thick line is composed by predicted 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.

vector of unknown parameters θ defining mean and heterogeneity in preferences. I adopt the following functional form:

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

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:

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

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

and Eqs. (3a) and (3b) describe household i's preference for housing characteristic m.[21] The term p_{ij}, which I call the user cost of the house in a given year, is defined as:

$$p_{ij} = rp_j + \tau_{ij} \quad (4)$$

where r is the annual interest rate. This simplified version of 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. To the best of my knowledge this is the first paper in this literature to allow for a specific housing price for each individual household.

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:

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

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:

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

which is increasing with income net of housing costs.

Given the household's problem described in Eqs. (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.,

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

where W_{ij} includes all of the non-idiosyncratic components of the indirect utility described in Eq. (2). As the inequalities in Eq. (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:

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

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

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

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

4.2. Endogeneity of housing prices

The main concern that arises in estimating MWTP in the framework of Eqs. (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 the characteristics of the house that affect utility. In other words, 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 et al. (1995) method.[22] I choose the control function approach because the high number of

[21] Unobserved heterogeneity is not modeled in this paper for two reasons. First, the micro-data allow the inclusion of a rich set of observed heterogeneity that gives rise to flexible substitution patterns. In practice, the empirical model has an assumption of no preference heterogeneity within those household demographic variables included in the estimation. Second, the standard method of including unobserved heterogeneity in choice models usually relies on an additional assumption about the underlying distribution of preferences (usually a normal distribution is used). This method is mainly used for aggregate data sets that are not able to include individual characteristics. Since this is not the case for this application, I preferred to rely on a simple structure and only control for observed heterogeneity. More generally, see Heckman (1997) and Angrist and Imbens (1999) for a discussion about the difficulties in interpreting, recovering, and extrapolating structural parameters when unobserved heterogeneity plays a critical role in individual decisions.

[22] Hausman (1978), Heckman (1978) and Smith and Blundell (1986) initially developed the control function, and it can be thought of as a two-stage least square approach applied to non-linear models. 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 et al. (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.

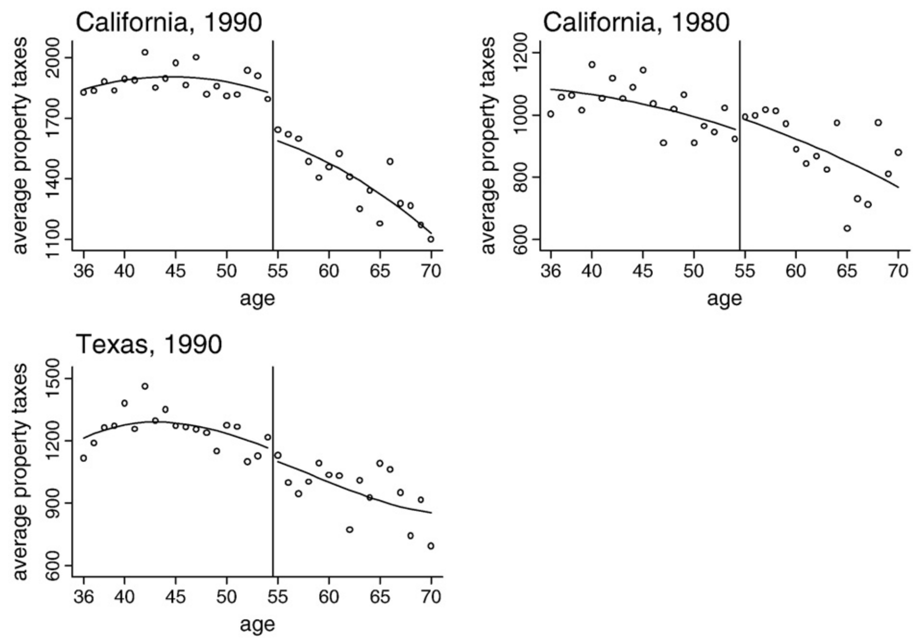

Fig. 8. Average property taxes for recent movers by age, California and Texas, 1990 and 1980.
Source: 1990 IPUMS. Notes: Each dot represents the average property taxes faced by each age group. The thick line is composed by predicted values of a polynomial regression of average property taxes on a dummy for 55-year, age, age squared and cubic, household characteristics and housing amenities.

products in the housing market leads to complications in estimating mean utilities for each product in a Barry et al. (1995) framework. Although both methods have slightly different properties–see Petrin and Train (2006)–they both face the same problem of finding credible instruments that are correlated with price and uncorrelated with the mean utility that all households share from each house.

The validity of the control function relies on an instrumental variable that directly affects housing prices, and that is not correlated with the error term in Eq. (2). The instrumental variable proposed here corresponds to the property tax component in the p_{ij} defined by Eq. (4), i.e. the component that varies according to Proposition 13 and its amendments. As we already estimated in Sections 2 and 3, property taxes in California are an important determinant of household mobility, and also correspond to a large share of the household income. As we will see in the empirical result, it also provides a strong first stage estimate.

I parametrically transform the characteristics of Propositions 13, 60 and 90 into an instrumental variable in the following way: For the chosen houses, the property tax is self reported in the Census data. For the alternative houses, the institutional framework of Propositions 13,

Table 3
RDD estimates of the effect of Propositions 60 and 90 on property taxes.

	Recent movers in California, 1990			Long tenure homeowners CA, 1990	Recent movers TX, 1990	Recent movers CA, 1980
Dependent variable: property taxes	(1)	(2)	(3)	(4)	(5)	(6)
1 if age ≥ 55	−0.08	−0.08	−0.08	−0.01	0.01	0.00
	−0.01	−0.02	−0.02	0.00	−0.06	−0.03
Age		−0.01	−0.01	−0.02	0.00	0.01
		−0.01	−0.01	0.00	−0.02	−0.01
Age controls (squared and cubic)	N	Y	Y	Y	Y	Y
Housing and household controls	N	Y	Y	Y	Y	Y
Metropolitan area dummies	N	N	Y	Y	Y	Y
R-squared	0.005	0.008	0.014	0.058	0.0535	0.018
Sample size	19,854	19,854	19,854	112,705	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.

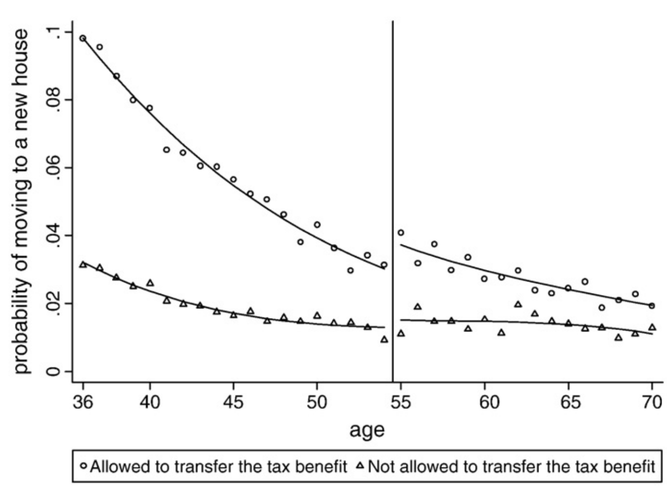

Fig. 9. Probability of moving by tax benefit transfer status, California 1990. Notes: The group of homeowners allowed to transfer the tax benefit is made by those homeowners who moved in 1989–1990 within the same county, or to a different county that approved Proposition 90. Each dot represents the average mobility rate by year faced by each group. 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.

60 and 90 is used to generate the taxation costs that a specific household faces when choosing any of those houses. For example, a 54-year old is assumed to have property taxes calculated at 1% of the house value for the alternative houses. 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; and b) the homeowner is moving within the same county or to a county that accepts Proposition 90.

In practice, this procedure embeds the Regression Discontinuity presented in Section 3 into an instrumental variable, generating a user cost of the house that is specific to each housing alternative and homeowner. With this constructed set of $\tilde{\tau}_{ij}$ I estimate the following first stage for the income effects specification, where the sample is a combination of chosen and alternative houses for each household:

$$g(I_i - \tilde{p}_{ij}) = \lambda\tilde{\tau}_{ij} + x_j\psi + \nu_{ij} \tag{10}$$

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

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

where γ_i also depends on observed household characteristics. As evident from Eq. (11), the predicted residual $\hat{\nu}_{ij}$ is a proxy for the unobserved housing quality ξ_j, with the additional benefit that $\hat{\nu}_{ij}$ is also specific to each household.[23] As in traditional two-stage least squares estimates, the identification strategy fundamentally relies on the first stage results.[24]

A pure Regression Discontinuity design is not used in the model for two reasons. First, the model is estimated with cross sectional data, and therefore there is no individual information about housing choices and property taxation before and after a household moves.[25] The framework above, instead, adapts the dynamic nature of those tax rules to the choice setup of chosen versus alternative houses. It does so by specifically incorporating the discontinuous variation in the ability to transfer the tax benefit due to Propositions 60 and 90 into the creation of the instrumental variable. Second, the initial variation in property tax payments generated by Proposition 13 is quite important, and identification in this model is not complete with just the variation from Propositions 60 and 90. These latter amendments provide variation in the ability to transfer the tax benefit, but Proposition 13 determines the magnitude of the tax break or the benefits lost in case of a household that moves.

Because of that, a potential concern about using $\tilde{\tau}_{ij}$ as a source of identification is that property taxes of the chosen house are correlated with homeowner tenure, i.e. the variation in $\tilde{\tau}_{ij}$ is not solely coming from differences in the ability to carry over a lower tax basis due to Propositions 60 and 90. If the property tax component of the user cost is correlated with length of stay in the same house, then it is also potentially correlated with unobserved housing quality, undermining the identification strategy. A simple regression of current property taxes on tenure produces an R^2 of 0.15. In order to address this issue, $\tilde{\tau}_{ij}$ is modified such that individual tenure is decomposed out of property taxes for the chosen house.[26] As expected, I find an R^2 of only 0.04 for the regression of the modified $\tilde{\tau}_{ij}$ on homeowner tenure. I use this adjusted $\tilde{\tau}_{ij}$ throughout the paper.

Finally, 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, the MWTP estimate for an extra room will be biased. To the best of my knowledge, only Bayer et al. (2007) deal with this issue to estimate valuation of school quality within a sorting model using school attendance zone boundaries. But the inclusion of possible different strategies for all estimated parameters is beyond the scope of this research.[27]

5. Estimation results

5.1. Data, choice set, and first stage

The empirical application uses 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[28]. 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 \$5000 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.

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

[23] Interactions of the predicted residual with choice variables were included in the model as a consistency check, leading to results relatively similar to the original specification.

[24] Since we are including the predicted unobserved component instead of the actual component, standard errors were adjusted by the following bootstrap procedure: a) estimate 1000 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 coefficients and add it to the standard errors obtained from the traditional formulas.

[25] If such information was available, one could specify that a household moves from old location j to a new location h only if $u_{ih} > u_{ij}$, therefore comparing changes in housing characteristics with changes in housing prices and property taxes for the same household, at ages 54 and 55. This setup also simplifies substitution patterns among alternatives and reduces the dimensionality problem of comparing current choice to all alternative houses in a given market. The decision process of each household reduces to the following comparison in utility flows (assuming a finite number of years $t+n$ and only one change in housing location): $u_{iht} + Bu_{iht+1} + B^2u_{iht+2} + \ldots + B^nu_{iht+n} > u_{ijt} + Bu_{ijt+1} + B^2u_{ijt+2} + \ldots + B^nu_{ijt+n}$. That in turn can be further adapted to include details of Propositions 13, 60 and 90 once the functional form in Eq. (2) is implemented. In practice, 54-year olds would have a constrained decision to move, since they are not allowed to transfer the tax benefit until they reach the age of 55, while households 55 years of age and older face an unconstrained decision. The magnitude of the constraint between those ages and their housing decisions would identify how much households value the new home relative to the old house.

[26] 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. Then, property taxes are estimated for the alternative houses using the institutional features of Proposition 13 and its amendments.

[27] Another potential confounder is that low property tax rates due to Proposition 13 acts as a commitment device that induces higher investments in social capital, and that high social capital could attract new homeowners that value that features, and therefore impact housing prices. However, the importance of such social capital is still a matter of debate in the literature.

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

Table 4
Household location decision model: raw preferences, California, 1990.

	Simple logit	Including covariates	Control function	Wealth effects	
				Logit	Control function
	(1)	(2)	(3)	(4)	(5)
User cost	−0.480	−0.797	−2.476	0.335	13.150
	0.005	0.010	0.017	0.009	0.065
Number of rooms	0.297	0.556	0.979	0.367	0.878
	0.005	0.008	0.010	0.008	0.009
1 if detached house	0.375	0.394	0.507	0.251	0.688
	0.004	0.010	0.011	0.010	0.011
Avg neighborhood income (/10 K)	0.147	0.276	0.856	−0.015	1.053
	0.005	0.008	0.012	0.007	0.011
% white	0.037	0.056	0.092	−0.112	0.101
	0.004	0.009	0.011	0.009	0.010
Control function			1.693		12.743
			0.007		0.062
Housing and neighborhood controls	N	Y	Y	Y	Y
Sample size	98,407	98,407	98,407	98,407	98,407

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 are bootstrapped, as reported in the text.

Table 5
Household location decision model: MWTP estimates, California, 1990.

	Simple logit	All variables	Control function	Wealth effects	
				Logit	Control function
	(1)	(2)	(3)	(4)	(5)
1 extra room	3372	3791	2151	33,701	2052
	(67)	(75)	(32)	(1107)	(24)
Detached house	19,797	12,515	5188	107,133	7484
	(333)	(364)	(117)	(5083)	(121)
Avg neighborhood income (/10 K)	960	1081	1082	−793	1415
	(32)	(36)	(27)	(391)	(16)
% white	3274	2985	1562	−80,005	1835
	(373)	(456)	(180)	(6633)	(175)
Housing and neighborhood controls	N	Y	Y	Y	Y
Average user cost	20,035	20,035	20,035	20,035	20,035
Sample size	98,407	98,407	98,407	98,407	98,407

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.

newly purchased in the previous year. This is the best proxy for houses available in the market in the year of analysis.[29] 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)–largely simplifying the dimensionality problem of comparing chosen house versus all other alternatives.

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.[30] Appendix Table A1 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.

After generating property taxes for all housing alternatives using the institutional features of Propositions 13, 60 and 90, the individual user cost of the house is constructed as in Eq. 8, using an interest rate of 6%.[31] Columns (2) and (3) of Appendix Table A2 report 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.

5.2. MWTP for housing amenities

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 four variables–two housing features: number of rooms and detached houses; and two neighborhood characteristics: average neighborhood income and percentage white in a block group–to compare how changes in the model affect 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 column (2), the estimated coefficient of the user cost variable remains relatively small in magnitude, even when including a broad set of housing and neighborhood controls.

Column (3) reports the estimate results for a specification similar to the one in column (2) but with the addition of a control function, equal to the residual of the first stage model for the user costs, as indicated in Eqs. (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 in column (1) shows very high MWTP estimates for those amenities: $3372 per year for an extra room, $19,797 for a detached house, $960 for a $10,000 higher average income of the neighborhood, and $3274 a marginal increase in the percentage white. Even after including other controls for housing and neighborhood amenities in column (2), the results still look very similar. Column (3) 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: $2151 for number of rooms, $5188 for a detached house, a surprisingly similar estimate for average income, and only $1562 for percentage white.[32]

Wealth effects are included in columns (4) and (5) of Tables 4 and 5. Again, the results are meaningful only after controlling for the

[29] Misspecification of the choice set may lead to 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. The choice set defined in this paper also does not include renters. This could be another source of bias if a significant fraction of homeowners consider renting as a credible housing alternative. The rent/own decision is not modeled in this paper since the instrumental variable proposed here does not apply to renters.

[30] 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 Arc View 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).

[31] Similar results were achieved with interest rates of 8% and 10%.

[32] Appendix Table A2, column (1) 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.

Table 6
Household location decision model: heterogeneity in MWTP estimates, California, 1990.

		MWTP		
	Baseline	Income (+$10,000)	Age (+10 years)	White (vs. other races)
	(1)	(2)	(3)	(4)
1 extra room	1783	1832	1801	1826
	(49)	(53)	(52)	(54)
Detached house	3509	3468	2339	3243
	(140)	(141)	(95)	(138)
Avg neighborhood income (/10 K)	1532	1574	1426	1523
	(36)	(38)	(35)	(37)
% white	2560	2590	2180	4690
	(280)	(290)	(260)	(490)
Housing and neighborhood controls	Y	Y	Y	Y
Average user cost	20,035	20,035	20,035	20,035
Sample size	98,407	98,407	98,407	98,407

Source: US Census Bureau — 1990 California Decennial Census Long Form data. Notes: Table shows MWTP estimates (and standard errors) derived from a multinomial logit model that includes heterogeneity, i.e., all choice variables were interacted with individual income, a dummy for white, and age. 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, first grade test scores, and distance to work. Standard errors were calculated using the delta method.

Table 7
Household location decision model: raw preferences and MWTP estimates, Dallas, Texas, 1990.

	Raw preferences		MWTP	
	Simple logit	Control function	Simple logit	Control function
	(1)	(2)	(3)	(4)
User cost	−0.155	−0.181		
	0.016	0.021		
1 extra room	0.078	0.093	2097	2137
	0.014	0.016	(431)	(447)
Detached house	0.066	0.066	10,658	9216
	0.011	0.011	(2135)	(1911)
Avg neighborhood income (/10 K)	−0.076	−0.071	−3333	−2665
	0.014	0.014	(705)	(625)
% white	0.018	0.019	5415	4950
	0.013	0.013	(3909)	(3384)
Control function		0.029		
		0.015		
Housing and neighborhood controls	Y	Y	Y	Y
Average user cost			9103	9103
Sample size	11,083	11,083	11,083	11,083

Source: 1990 IPUMS. 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. Standard errors for the control function approach are adjusted by the following bootstrap procedure: a) estimate 1000 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 coefficients and add it to the standard errors obtained from the traditional formulas.

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. As expected, the MWTP estimates are usually higher for all variables when the model includes wealth effects. Homeowners are on average willing to pay a similar amount for an extra room, but 45% more for a detached house, 30% more for a higher average income, and 18% more for additional whites in the neighborhood.

Table 6 reports parameter estimates for a model that includes observed heterogeneity. 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 the baseline estimates are not comparable to the results from the homogeneous model on Table 5, they are reasonably similar. 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 very strong preferences to live with individuals of the same type.

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.[33] 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 shown in column (2). The control function coefficient estimate is not significantly different from zero, and its magnitude is very small (0.029) 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 control function does not provide new relevant variation to the estimates. Given that house values in Texas are re-assessed every 2 to 3 years, moving costs represented by property taxes do not carry much information about historical house prices, and potential house buyers have similar user costs for the same property in that state.

6. Conclusion

Given that homeownership is the primary method in which families accumulate wealth, understanding housing demand is of special importance in evaluating questions of welfare and equity across household types. Unfortunately, most 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. Due to its importance, in fact, recent empirical and theoretical research has incorporated moving costs, such as Bayer et al. (2009) and Epple et al. (2009).

In this paper I provided clear evidence that transaction costs affect individual behavior. Using a natural experiment 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. In fact, 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.

The research design developed in this paper adapted a dynamic problem related to property taxation cost to a cross-sectional housing choice model. Future work may directly incorporate transaction costs, and estimate MWTP by comparing housing decisions of the same household over time. Although such task is certainly dependent on improving current data sets used to study housing markets, it would allow more flexible assumptions about choice sets and unobserved heterogeneity.

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[33] 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.

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Appendix Table A1. Average household characteristics, California 1990.

	30 to 70-year old homeowners	Choice set
	(1)	(2)
House value	306,984	301,609
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	64,546	63,941
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	98,407	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 A2. Hedonic regression and first stage estimates, California and Texas, 1990.

	California			Texas
	Hedonic regression	Control function	Wealth effect	Control function
Dependent variable: user cost	(1)	(2)	(3)	(4)
Adjusted property taxes		4.0	−0.0003	4.1
		(0.004)	(0.000003)	(0.03)
Number of rooms	1549	792	−0.022	605
	(13)	(3)	(0.003)	(21)
1 if detached house	2434	1916	−0.146	−115
	(60)	(14)	(0.011)	(108)
Avg neighborhood income (/10 K)	1327	599	−0.032	685
	(9)	(3)	(0.002)	(35)
% white	4843	1657	−0.126	−1545
	(115)	(31)	(0.024)	(220)
Housing and neighborhood controls	Y	Y	Y	Y
R-squared	0.57	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.

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