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Information, Search, and House Prices: Revisited

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Abstract Buyers pay different prices for nearly identical homes. One explanation for this is that housing markets are thin, resulting in price bargaining between sellers and buyers. If the relative bargaining power of buyers varies, so will sales prices. One hypothesis is that the relative bargaining strength of buyers coming from outside the local market relative to that of local residents is weak, because distant buyers have high search costs and may know less about the nuances of the local market. Our results, based upon a large number of single-family home transactions from the state of Florida, lend support to this hypothesis. Another related hypothesis is that buyers' price expectations are anchored to prices they were accustomed to at their previous residence. Hence, if they come from high price markets they will tend to pay more for their new home. This hypothesis is also supported by our results.

Keywords Housing prices · Search · Information · Anchoring

Introduction

A great deal of research shows that housing prices depend on the characteristics of the housing bundle, broadly defined to include structural characteristics, neighborhood characteristics, location, and public services. But research also shows that controlling for these characteristics, there is not one price but rather a distribution of prices within the local housing market. One explanation for the housing market's violation of the "law of one price" is that in thin markets, prices are negotiated and

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therefore the price is affected by the relative bargaining power of sellers and buyers. In this context, a number of studies have shown that the racial and demographic characteristics of sellers and buyers affect the sales prices of similar housing units (Harding et al. 2003; Ihlanfeldt and Mayock 2009a).

Apart from the influence of racial and demographic characteristics, it has been hypothesized that the relative bargaining position of homebuyers is weaker if their previous residence was located outside, rather than within, the local housing market. Turnbull and Sirmans (1993) offer two explanations for why in-migrants may pay more for housing than local residents: (1) due to travel expenses, in-migrants may have higher search costs, and (2) because they do not have first-hand experience observing unique market conditions over a longer period of time, in-migrants may be less knowledgeable about the local distribution of housing prices than local residents. If local residents perceive the mean price of housing accurately, for instance, whereas in-migrants have a perceived mean price that exceeds the actual mean, then in-migrants will pay more for housing. Lambson et al. (2004) suggest that one reason in-migrants may have prior beliefs biased on the high side is that they may come from high-price housing markets. If distant buyers do pay more for housing than local residents, this has implications for market efficiency, because it suggests that the Multiple Listing Service (MLS) and other housing market institutions do not provide buyers with sufficient market information to ensure that different types of buyers pay similar prices for comparable units.

Three papers provide evidence on the hypothesis that distant buyers, in comparison to local residents, are at a disadvantage within the local housing market. The results of two of the studies fail to support this hypothesis (Turnbull and Sirmans 1993; Watkins 1998), while the evidence provided by the most recent study (Lambson et al. 2004) strongly supports this hypothesis. Hence, there is disagreement over whether the housing market is efficient in the sense that existing institutions effectively transmit sufficient information that the distribution of prices for comparable homes does not systematically differ across different types of buyers. Because of this dissension and the fact that the data used by all of these studies have significant limitations, there is a clear need for a more convincing test of the distant buyer hypothesis.

In this paper we make use of new data from the state of Florida that provides both the current and prior residential locations of a sample of approximately 7,000 homebuyers. These data allow us to overcome the methodological weaknesses of prior studies that have tested the distant buyer hypothesis.

Using two different distance measures, the results lend strong support to the hypothesis that non-local residents pay more for housing. Using a wide variety of alternative tests, we also find that buyers pay more if they come from a market where prices are higher. The latter result is consistent with the "anchoring" hypothesis which argues that buyers' price expectations are anchored to what they have grown accustomed to in their previous residence.

Literature Review

Turnbull and Sirmans is the first study to provide evidence on whether out-of-town buyers pay more for comparable housing than buyers who reside within the market area. Their data consist of 151 single-family home sales that occurred over a one-year time period from similar subdivisions located in Baton Rouge, Louisiana. A standard hedonic price model was estimated that included a dummy variable indicating whether the buyer was new to the Baton Rouge metropolitan statistical area. The estimated coefficient on this dummy variable, while positive as expected, was not statistically significant. The authors concluded that the housing market is sufficiently efficient to not allow different prices for comparable homes independent of the amount of information that buyers possess.[1]

The results of Turnbull and Sirmans are questioned by Lambson et al., who suggest that the Turnbull and Sirmans test has little power since the Turnbull and Sirmans sample only included 63 non-resident buyers. Lambson et al. also suggest that Turnbull and Sirmans's results may have been biased by omitted variables because their models contained too few covariates. The Turnbull and Sirmans model includes living area and "other area" as property descriptors.

Watkins (1998) estimates a hedonic equation similar to that of Turnbull and Sirmans in an attempt to replicate their study using sales transactions from Glasgow in the United Kingdom. In comparison to Turnbull and Sirmans, Watkins' sample is larger (544 total observations with 138 non-residents), and he includes a more extensive set of property descriptors. His results mirror those obtained by Turnbull and Sirmans, with no difference in price found between intra-market movers and in-migrants. However, the criticisms that Lambson et al. make of Turnbull and Sirmans can also be applied to Watkins' study, given that his non-resident sample is still small and his list of descriptors fails to adequately control for location-specific amenities known to affect property values.[2]

The final study by Lambson et al. differs from the two previous studies in many respects. First, Lambson et al. focus on the market for multi-family structures rather than the single-family home market. Their data set includes 2,854 apartment building transactions that occurred in the Phoenix metropolitan area from 1990 to 2002. Second, instead of defining a distant buyer as someone originally living in another city, they define a distant buyer as someone from a state other than Arizona. Third, unlike the previous two studies that found no difference in the prices paid among different types of buyers, Lambson et al. find that non-Arizona residents pay a premium of about 5.5% in comparison to within-Arizona buyers. Finally, Lambson et al. examine whether the distant buyer premium is attributable to a search cost disadvantage or whether it is due to out-of-state buyers having inflated perceptions of Phoenix's price distribution as the result of coming from higher-price markets.[3] Their results provide weak evidence that distant buyers have a search cost disadvantage and that they are anchored to higher prices.

[1] Their conclusion was also based on their finding that there was no price difference between first-time versus repeat buyers.

[2] For example, there is no variable accounting for differences in perceived public school quality, which surely varies within a city of Glasgow's size and has been shown to have a strong impact on housing price. For educational continuity, the moves of local residents may be constrained to market areas that would allow their children to remain in their current schools, even if these schools are of lower quality. In-migrants, on the other hand, have no such constraint and may therefore end up buying in better school zones.

[3] Lambson et al. refer to the latter possibility the "anchoring hypothesis" based on the idea that buyers' beliefs are anchored on what they have become accustomed to in the past.

Theoretical Framework

The search theoretic model underlying our empirical work has already been formalized by Turnbull and Sirmans and Lambson et al. and therefore need not be repeated here. However, the basic results can be simply expressed. Because the time and monetary costs associated with shopping for a new home generally increase with the distance between a buyer's current home and prospective neighborhood, buyers making out-of-town moves will generally have higher search costs than buyers making in-town moves. Moreover, out-of-town buyers may face a binding constraint on the length of search.[4] Both of these differences place out-of-town buyers at a bargaining disadvantage, and out-of-town buyers are subsequently expected to pay more than in-town movers for a home of similar quality.

A buyer's relative bargaining power is also influenced by his knowledge of the local house price distribution. For example, if a buyer overestimates the mean of the price distribution, it can be shown that he will be willing to pay more for a home than buyers who know the true mean.[5] Both Turnbull and Sirmans and Lambson et al. argue that distant buyers may have relatively poor information of local prices because locals know more about location-specific information such as trends in growth, zoning, crime, and so forth. Hence, the gap between the perceived and actual mean may be greater for distant buyers. On average, the difference between the perceived and actual mean price is expected to be positive, resulting in distant buyers paying more. Because realtor commissions are a percentage of the sales price, if buyers underestimate the mean price, realtors have the incentive to correct their misperceptions. Moreover, Lambson et al. emphasize that price beliefs are affected by anchors, with one such anchor being the prices the buyer was accustomed to within his previous market. Hence, if distant buyers tend to come from higher price markets, they are more likely than local movers to overestimate the mean of the housing price distribution. Such buyers would subsequently be willing to pay higher prices when bargaining for a house and would, on average, pay more for housing.

The aforementioned theoretical results imply that different types of buyers will pay different prices *for an identical home*. As many of the characteristics of housing valued by consumers are unobserved to econometricians working with even the most detailed of datasets, estimating the effects of bargaining characteristics is complicated by the potential bias from such omitted variables.[6] For instance, if out-of-town buyers, or buyers from higher-priced neighborhoods, purchase homes of higher (unobserved) quality, the estimated coefficients on variables measuring these buyer attributes would be biased upward in a standard hedonic model. Identifying the effects of these buyer characteristics subsequently necessitates both a detailed dataset as well as an estimation strategy that mitigate the influence of unobserved housing characteristics on the buyer-type coefficients in the regression equation, which we now describe in turn.

[4] For example, out-of-town buyers with school-age children will generally want to move into their new residence before a new school year begins. In-town movers, particularly those moving within the same school district, will not face such constraints on the search process.

[5] See Turnbull and Sirmans (p. 548) for the derivation of this result.

[6] This issue is investigated extensively by Harding et al. (2003), who suggest an innovative identification strategy for addressing the omitted variables problem.

Data

Our data come from the standardized property tax rolls that each of Florida's 67 counties must submit to the Florida Department of Revenue (FDOR) annually for auditing purposes. From these rolls we extracted all recent sales of single-family homes located within Florida's metropolitan areas. The information contained on these rolls changed in 2009 as the result of Florida passing an amendment (Amendment One) to its constitution in 2008 providing property tax relief to Florida's citizens. One provision of this amendment is built upon an earlier amendment drafted in 1992 ("Save Our Homes") that limited annual increases in the assessed value of property receiving the homestead exemption to 3% or the percentage change in the Consumer Price Index, whichever is lower. Prior to Amendment One, tax savings resulting from a difference between market and assessed values were lost if the homeowner moved. For the new home, the assessed value was set at the home's market value. Amendment One allowed homeowners to carry their tax savings from their old home to their new home, a provision called portability. To implement portability the 2009 tax rolls included the parcel identification number of the previous residence of 2007 and 2008 homebuyers. Using parcel identification maps provided by the county property appraisers to the FDOR, we are able to determine the exact location of both the homebuyers' current and previous residences.[7]

In addition to current and previous locations, the data include information on the sales price (SALE), an identifier indicating whether the sale is arm's-length, and a number of property descriptors such as the age of the home (AGE) and the amount of interior living space (TOTLIV). We recognize that this list of descriptors may be too limited to avoid omitted variables bias in our model explaining sales prices. However, the tax roll data also provide the county property tax assessor's estimate of market value (MVALUE). Including the MVALUE as an explanatory variable in our sales price regressions obviates the need to include any property descriptors (although we do so for reasons to be outlined below), since MVALUE summarizes into a single number the locational and structural characteristics of the property.[8] MVALUE is estimated by the assessor to reflect the market value of the property as of January 1 of the tax roll year.[9] Sales are observed after this date; hence, MVALUE is independent of the actual sales price of the unit. In prior work (Ihlanfeldt and Mayock 2009a), we have found that 1) MVALUE is a good predictor of sales price, and 2) in comparison to the property descriptors, MVALUE is far more successful in explaining differences in housing quality across housing units.

As the current and previous residences of the buyer of the property are known, we can measure differences in search costs (defined here to also capture differences among buyers in their knowledge of the local housing market) in several different

[7] As the identifier for the previous residence is only reported for buyers that have made use of portability, our sample is restricted to intra-Florida movers who previously owned a home in the state.

[8] The use of tax assessors' estimates of market value in lieu of including property characteristics in a hedonic price model was pioneered by Clapp and Giaccotto (1992).

[9] County property appraisers rely on Computer Assisted Mass Appraisal (CAMA) programs that use information from recent sales to predict what each property would sell for on January 1 of the tax roll year. The Florida Department of Revenue annually audits each county's market value estimates for accuracy.

ways. Following the extant literature, our first measure of search costs is a simple indicator variable (NEWMARKET) that assumes a value of unity if the buyer made an inter-market move and zero if the move was intra-market. In the construction of this variable, we use the metropolitan statistical area to define distinct housing markets, and a move is considered inter-market if the buyer comes from outside the metropolitan area; either from a non-metropolitan area or another metropolitan area.[10]

Our second measure of the costs of search is the simple straight-line distance in miles between the owner's current and previous residences (MILESFROMHOME). Whereas use of the NEWMARKET variable implicitly assumes that search costs are uniform for all movers coming from outside of the market (but may differ from local movers, who are assumed to have their own uniform level of search costs), specifications in which MILESFROMHOME serves as the search cost measure allow for such costs to vary smoothly with the length of the buyer's move, regardless of whether it originates from within or outside the local metropolitan area. Allowing for such a relationship may be important given the size of the state of Florida and the possibility that search costs may increase monotonically with distance.[11]

The information about the buyer's previous residence is also utilized to generate the variable we use to test the anchoring hypothesis. Recall, this hypothesis states that buyers' bids are affected by the home prices they have been accustomed to at their previous residence. After locating each buyer's previous location on a digital map provided by FDOR, this location is then assigned to a county, jurisdiction, and census block group. Using the information on the year and month in which the buyer purchased the new property, we calculate the mean price per square foot for all sales occurring in the previous year in the buyer's previous housing market. This mean is calculated using three different definitions of the buyer's home market: the block group, the jurisdiction, and the county. These variables are labeled PREVIOUSPFOOT_BG, PREVIOUSPFOOT_JURIS, and PREVIOUSPFOOT_COUNTY. If buyers utilize recent sales activity in the housing markets with which they are the most familiar to estimate the price distribution during their search, then these anchoring variables proxy for differences among buyers in their perceptions of the price distribution.

As is well known, in recent years Florida's housing markets have been highly volatile. This volatility created an issue in our use of the data. Though our sample is restricted to arm's-length transactions of single-family homes, conversations with FDOR employees revealed that county property appraisers in a number of Florida counties were classifying distressed sales (e.g., foreclosures and short sales) as arm's-length transactions. As the current FDOR data do not contain any information on whether or not a sale is distressed, including such sales and failing to control for distressed conditions could introduce substantial omitted variables bias into our estimated sales price regressions. To guard against such bias, we utilize information

[10] We utilize the U.S. Office of Management and Budget's November 2008 definitions for metropolitan statistical areas, accessible at the following URL: http://www.census.gov/population/www/metroareas/metrodef.html. Sales occurring in non-metropolitan areas are excluded from the analysis when using the NEWMARKET variable, as it is unclear how housing markets should be defined in non-metropolitan areas.

[11] Such a relationship may arise, for instance, if transportation costs increase substantially with distance.

from the Department of Housing and Urban Development's (HUD) Neighborhood Stabilization Program data. Created to help local governments identify neighborhoods in danger of becoming blighted, this data set contains an estimated foreclosure rate for each census tract throughout the nation.[12] An inspection of the HUD data suggested that census tracts with foreclosure rates in excess of 8% were in the tail of the distribution; subsequently, sales in these census tracts were dropped from the sample to guard against the possibility of the omitted variables bias associated with the inclusion of distressed sales in the sample.[13]

The final data set is comprised of 6,666 single-family home sales. Of these sales, 977 buyers are making inter-market moves. Sample means and standard deviations for the variables used in our empirical model are reported in Table 1. While our sample size is large relative to those used in previous studies of the single-family market, we are only able to study homeowners moving within the state of Florida. However, the importance of this limitation is mitigated by Florida's large size and the diversity of its housing markets.

Econometric Framework

To estimate the effect of search costs and anchoring effects on the recorded transaction price, we estimate the following regression model

$$\begin{aligned}\ln(SALE_i) = &\sum_{\mathrm{j=1}}^{\mathrm{K}} \mathrm{C_j} \ln(MVALUE_i)\beta_{1,j} + \ln(TOTLIV_i)\beta_2 \\ &+ \ln(AGE)\beta_3 + \ln(PREVIOUSPFOOT_i)\beta_4 \\ &+ DISTANCE_i\beta_5 + \sum_{j=1}^{K} \mathrm{C_j}\tau_j\beta_{6,j} + \sum_{t=1}^{T} M_t\beta_{7,j} + \varepsilon_i\end{aligned} \tag{1}$$

where $\mathrm{C_j}$ denotes a series of indicator variables that assume a value of one if the sale occurred in county j; τ_j denotes a county-specific trend; and M_t denotes a series of dummy variables assuming a value of one if sale i occurred in month t. $DISTANCE_i$ represents one of our two measures of search costs (ln(MILESFROMHOME) or NEWMARKET). Equation 1 allows for differing relationships between the purchase price and estimated market value across counties, county-specific price trends, and macroeconomic shocks that influence property values throughout the state.

While MVALUE may be a good predictor of a unit's sales price, it remains an estimate and is therefore subject to measurement error. If this error is correlated with buyer characteristics, our estimate of the distance premium may be biased. To the extent that measurement errors are correlated with property descriptors, and the latter are correlated with buyer characteristics, the potential for biased estimates of the distance premium is mitigated by including the age and living space variables (ln (AGE) and ln(TOTLIV)), along with the estimated market value (ln(MVALUE)), in the model.

[12] More information on the HUD Neighborhood Stabilization Data can be found at http://www.huduser.org/portal/datasets/NSP.html.

[13] Of the 2,000 census tracts included in our sample 685 had estimated foreclosure rates in excess of 8%. Dropping sales in these tracts eliminated 2,006 sales from the sample.

Table 1 Summary statistics

Variable	Description	Sample		
		All movers[a]	Intra-market movers[a]	Inter-market movers[a]
		Mean	Mean	Mean
SALE	Recorded sales price of property	345909.6	350738.4	318204.8
		[173360.3]	[176020.1]	[154170.8]
MVALUE	Assessor's estimate of property's market value on January 1st in the year-of-sale	328086.7	334438.4	290707.0
		[171873.7]	[175911.4]	[139707.0]
TOTLIV	Interior living space of primary improvement on the parcel	2689.9	2714.4	2547.4
		[901.3]	[903.5]	[875.0]
AGE	Age of the primary structure on the parcel	13.9	14.2	12.5
		[11.6]	[11.7]	[10.8]
NEWMARKET	Dummy variable equal to one if buyer moved from outside of the metropolitan area	0.15	–	–
		[0.35]	–	–
MILESFROMHOME	Distance between the buyer's current home and the buyer's previous home	23.7	6.6	123.7
		[51.9]	[8.0]	[79.4]
PREVIOUSPFOOT_BG	Mean SALE/TOTLIV: Properties sold in previous year in buyer's previous Census Block Group	139.7	139.3	142.3
		[47.4]	[46.4]	[53.1]
PREVIOUSPFOOT_JURIS	Mean SALE/TOTLIV: Properties sold in previous year in buyer's previous jurisdiction	134.8	134.7	135.7
		[37.0]	[36.2]	[41.0]
PREVIOUSPFOOT_COUNTY	Mean SALE/TOTLIV: Properties sold in previous year in buyer's previous county	134.2	133.9	135.9
		[30.6]	[29.4]	[36.8]
Observations		6,666	5,689	977

Standard deviations in brackets

[a] Sample does not contain sales from non-metropolitan counties

Our final strategy for combating bias resulting from measurement error in the assessor's market value estimate is to include fixed effects for the buyer's neighborhood unit. In effect, our models compare homes purchased by distant versus local buyers within the same neighborhood that have the same estimated market value. Even if measurement errors are correlated with the mix of distant versus local buyers across neighborhoods, as long as these errors are uncorrelated with the type of buyer within neighborhoods, the inclusion of the fixed effects will result in unbiased estimates. To define our neighborhood unit, we again draw upon the FDOR tax roll data. Appearing on these tax rolls are variables indicating the township, range, and section (TRS) from the Public Land Survey System (PLSS) within which the parcel is situated.[14] The TRS can be used to place each parcel within a one-mile by one-mile square area.[15]

Results

The estimated parameters from Eq. 1 are reported in Table 2. Reading from left to right, the first three columns report the estimated parameters and standard errors (clustered at the TRS level) associated with TRS fixed effects models (fixed effects models hereafter), where search costs are measured using the natural logarithm distance between a buyer's current and previous property. Each of these models includes a measure of the natural logarithm of the price per square foot of sales in the buyer's previous market, where the market definition varies across the specifications.[16] The next three columns present the results from estimating similar models where the measure of search costs is the NEWMARKET variable.

In each of the six fixed effects models, the empirical results are strongly in accord with theory, as more distant buyers pay statistically significant price premia for housing using both measures of search costs. The estimated coefficient on the NEWMARKET variable in Columns 4 through 6, for instance, implies that buyers switching housing markets pay approximately 1.9% more for identical housing than buyers making intra-market moves. For the mean sales price in the sample, this corresponds to a premium of approximately $6,600. This premium is substantially smaller than the 5.5% out-of-state premium estimated by Lambson et al. in the multi-family market. However, our results provide support for the distant-buyer hypothesis as it applies to owner-occupied housing and, therefore, stand in stark contrast to the

[14] More information on the PLSS can be found at http://nationalatlas.gov/articles/boundaries/a_plss.html. Each township/range/section combination typically corresponds to a one-mile by one-mile square. Although quite rare, it can be the case that the TRS code identifies a geographic area that is larger or smaller than one square mile. This might be the case, for instance, in sections near jagged borders (e.g., near the coast).

[15] As a basis for comparison, TRS squares are generally smaller than census tracts. In suburban Miami-Dade County, for instance, there are approximately four TRS squares in a typical census tract.

[16] All of the models reported here express the natural logarithm of the sales price as a function of the natural logarithm of each of the regressors. In another set of models, we estimated equations in which the natural logarithm of the sales price was expressed as a function of the test variables expressed in levels. The results from estimating these log-linear models were qualitatively similar to the results from the log-log specification. As both classes of models appeared to fit the data equally well, we focus here on the log-log results because the coefficients can be interpreted as elasticities.

Table 2 Fixed effects regressions: baseline specification

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
ln(MILESFROMHOME)	0.00322**	0.00304**	0.00312**			
	(0.00147)	(0.00149)	(0.00150)			
ln(PREVIOUSPFOOT_BG)	0.0535***			0.0527***		
	(0.0110)			(0.0110)		
ln(PREVIOUSPFOOT_JURIS)		0.0368***			0.0367***	
		(0.0139)			(0.0137)	
ln(PREVIOUSPFOOT_COUNTY)			0.0206			0.0214
			(0.0170)			(0.0168)
ln(TOTLIV)	0.278***	0.276***	0.276***	0.277***	0.275***	0.275***
	(0.0203)	(0.0205)	(0.0205)	(0.0204)	(0.0205)	(0.0205)
ln(AGE)	−0.0191***	−0.0192***	−0.0192***	−0.0194***	−0.0195***	−0.0195***
	(0.00444)	(0.00445)	(0.00444)	(0.00445)	(0.00445)	(0.00444)
NEWMARKET				0.0188***	0.0190***	0.0194***
				(0.00660)	(0.00656)	(0.00656)
Observations	6,666	6,666	6,666	6,666	6,666	6,666
R-squared	0.810	0.809	0.809	0.810	0.809	0.809
Number of neighborhood fixed effects	2,847	2,847	2,847	2,847	2,847	2,847
TRS fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

statistically insignificant distance premia estimated by the two previous studies (Turnbull and Sirmans 1993 and Watkins 1998).[17]

The coefficient on our second measure of search costs (ln(MILESFROMHOME)), which allows for search costs to increase continuously with the distance between the new and previous home, is also positive and statistically significant. As with our other distance measure, the magnitude of the ln(MILESFROMHOME) coefficient is economically significant. A one standard deviation increase from the mean in MILESFROMHOME (an increase of approximately 220%) increases the price of the mean-valued property by approximately $2,440.

Turning now to the results relevant to the anchoring hypothesis, when the priors of the buyer are measured using the price per square foot of homes sold within the block group or the jurisdiction, our results provide evidence in support of the hypothesis: the coefficient on the ln(PREVIOUSPFOOT_BG) and ln(PREVIOUSPFOOT_JURIS) variables are positive and statistically significant in each of the fixed effects models. A one-percent increase in PREVIOUSPFOOT_BG is associated with an increase in the sales price of the home of approximately 0.05%; the comparable elasticity using PREVIOUSPFOOT_JURIS as the anchoring variable is approximately 0.04. Using the estimated coefficients on ln(PREVIOUSPFOOT_BG), a one standard deviation increase in PREVIOUSPFOOT_BG from the mean (an increase of approximately 34%) results in a $6,278 increase in the price of the mean-valued property. A comparable increase in the price per square foot when the anchoring variable is constructed at the jurisdiction level corresponds to a price premium of $3,494.

It is important to note that regardless of how we control for search cost differentials, the magnitudes of the anchoring coefficients are monotonically declining as we enlarge the geographic area of the buyer's previous market from the block group to the jurisdiction to the county. The county-level anchoring coefficients are not only relatively small, but they are also statistically insignificant in both specifications (Column 3 and Column 6). There are a number of alternative explanations for these results, which we consider in the following paragraphs.

If the block group is the true geography at which a buyer's priors about the price distribution are formulated, then calculating the anchoring variable at the jurisdiction and county levels (geographies generally much larger than block groups) would

[17] In his discussant's comments, Geoffrey Turnbull offered an alternative explanation for our results; namely, that they are due to local buyers more frequently purchasing "for-sale-by-owner" (FSBO) homes. The validity of this explanation hinges upon whether 1) local buyers do in fact more frequently purchase FSBO homes, and 2) FSBO homes, *ceteris paribus*, sell for less than homes listed with a broker. We could find no evidence on 1), but it does have intuitive appeal, since it is reasonable to believe that local buyers are under less time pressure to find a new home, and therefore can visit a larger number of prospective neighborhoods and discover more FSBO homes. Regarding 2), the evidence is highly mixed. In their review of the studies that have tested the hypothesis that broker-listed homes sell for more, Yavas and Colwell (1995) cite 3 studies whose results support the hypothesis and an equal number whose results did not support the hypothesis. Their own evidence also did not support the hypothesis; in fact, they found that the seller's use of a multiple listing service reduced the sales price by 5.7%. The intuition underlying this result is that brokers search less when a the seller quotes a higher price because a higher price reduces the probability that a buyer will purchase the house. The most recent evidence on the hypothesis comes from Hendel et al. (2009), who find that listing the house shortens the time it takes to sell the house, but does not increase the final sales price.

introduce substantial measurement error, attenuating the anchoring coefficient. The empirical results reported in Table 2 are consistent with such attenuation.

A second explanation for our anchoring results is that buyers moving from high-price block groups may purchase homes that have more unobservable amenities than buyers moving from low-price block groups. Should this be the case, then the coefficient on ln(PREVIOUSPFOOT_BG) would register both this demand effect as well as the anchoring effect, which would account for its relatively larger magnitude. We guard against the possibility of a demand effect contaminating our estimate of the anchoring effect by including neighborhood-level fixed effects defined over small areas. However, we cannot rule out the possibility, a priori, that buyers moving from high-priced neighborhoods purchase homes that, *within the receiving neighborhood*, have more unobservables that positively influence the sales price of the home. This possibility is further investigated below.

A final explanation for the difference in the anchoring coefficients across models is that the changing magnitudes are simply a statistical artifact. As we calculate the anchoring variable at progressively larger levels of geography, we are reducing variance in the anchoring measure by smoothing the price per square foot variable over more and more space. If such smoothing reduces the within-neighborhood variance in the anchoring terms, then we would expect that it would become harder to identify the influence of anchoring working with progressively larger geographies.

The above explanations have a number of implications. First, as suggested by our second explanation, the anchoring variables may be correlated with unobservable demand effects, resulting in omitted variables bias. Omitted variables bias may also result from correlations between the distance variables and unobservables. We attempt to mitigate these potential omitted variables biases by including neighborhood fixed effects for small areas. Because fixed effects were not used in previous studies, it is of interest to re-estimate our models excluding the fixed effects.

A comparison of the estimated models including fixed effects (Table 2) with those that do not (Table 3) reveals a number of important differences in the results. Starting with the search cost variables, in the ordinary least squares (OLS) models the distance measure is statistically significant in none of the six models. Compared with the distance coefficients from the fixed effects models, the NEWMARKET coefficients are also significantly smaller in the OLS models, and the coefficients on ln(MILESFROMHOME) are negative and very small. The anchoring variables also change substantially between the OLS and fixed effects models: the PREVIOUSPFOOT coefficients from the OLS models are substantially larger than those of the fixed effects models, likely reflecting the large amount of heterogeneity in the housing stock in different neighborhoods throughout the state of Florida.

The differences between the two sets of models suggest that neighborhood-level unobservables are correlated with buyer characteristics and that this correlation imparts substantial bias onto the search cost and anchoring coefficients. That this bias persists even with our inclusion of the appraiser's lagged estimate of market value suggests that previous studies, which controlled less well for other factors affecting market value, may have been subject to substantial omitted variables bias. Such bias may explain the mixed results in the existing literature.

Table 3 Ordinary least squares regressions: baseline specification

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
ln(MILESFROMHOME)	−0.000127	−0.000422	−0.000283			
	(0.00126)	(0.00127)	(0.00129)			
ln(PREVIOUSPFOOT_BG)	0.0977***			0.0971***		
	(0.00943)			(0.00945)		
ln(PREVIOUSPFOOT_JURIS)		0.0723***			0.0709***	
		(0.0120)			(0.0120)	
ln(PREVIOUSPFOOT_COUNTY)			0.0295*			0.0259*
			(0.0151)			(0.0150)
ln(TOTLIV)	0.128***	0.113***	0.104***	0.128***	0.113***	0.104***
	(0.0130)	(0.0129)	(0.0130)	(0.0130)	(0.0129)	(0.0131)
ln(AGE)	0.00411*	0.00525**	0.00583**	0.00422*	0.00544**	0.00599**
	(0.00229)	(0.00233)	(0.00233)	(0.00229)	(0.00233)	(0.00233)
NEWMARKET				0.00569	0.00721	0.00869
				(0.00558)	(0.00556)	(0.00555)
Observations	6,666	6,666	6,666	6,666	6,666	6,666
R-squared	0.894	0.893	0.892	0.894	0.893	0.892
Number of neighborhood fixed effects	NA	NA	NA	NA	NA	NA
TRS fixed effects?	No	No	No	No	No	No
County-specific trend?	Yes	Yes	Yes	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Sensitivity Tests

Because theory provides little more than qualitative guidance about the relationships between distance, price anchors, and transaction prices, we estimated a number of additional fixed effects regression equations utilizing alternative functional forms and measures of the price anchors. To estimate our first alternative specification, we calculate the mean price per square foot for all sales occurring in the market into which the buyer moved (CURRENTPFOOT). As in our baseline specification, this mean is calculated using three different definitions of the buyer's home market: the block group, the jurisdiction, and the county. For each of the three geographies, we then calculate the natural logarithm of the ratio of PREVIOUSPFOOT to CURRENTPFOOT. These variables are labeled ln(PFOOTRATIO_BG), ln(PFOOTRATIO_JURIS), and ln(PFOOTRATIO_COUNTY) and are entered into Eq. 1 in lieu of the ln(PREVIOUSPFOOT) terms included in our baseline specifications. Results from the regressions including these ratio terms are reported in Table 4.

There are two important differences in the assumptions underlying our baseline specification (Table 2) and our first alternative specification (Table 4). First, the models in Table 2 assume that anchoring effects are independent of the characteristics of the market into which the buyer is moving. For example, in these specifications, even if a buyer is moving within the same block group, higher prices within that block group are assumed to increase the buyer's estimate of the mean of the price distribution. In the models in Table 4, however, anchoring is only permitted to affect buyer behavior *if the buyer's price anchors from his previous market differ from the price anchors in the market into which he is buying.*[18] The specifications reported in Table 4 also differ from those in Table 2 in that the Table 4 equations impose the assumption that the influence of price anchors is relative, not absolute. For instance, if a buyer moved from a high-priced market to another high-priced market, the specifications in Table 4 allow for the anchoring effect to be mitigated by the fact that the buyer's current and previous market are similar; the specification in Table 2 does not allow for such an effect. As in the Table 2 models, positive coefficients on the ln(PFOOTRATIO) terms reported in Table 4 are interpreted as support for the anchoring hypothesis.

The results using the ratio variables show that the estimated distance effects are generally insensitive to the change in the anchoring variables, as the coefficients on the NEWMARKET and ln(MILESFROMHOME) variables in Table 4 are very similar in sign and statistical significance to those reported in Table 2. Moving on to the anchoring coefficients, the results in Table 4 are consistent with those reported in Table 2: 1) the anchoring hypothesis is supported under the block group and jurisdiction definition of the anchoring geography, and 2) the estimated magnitude and statistical significance of the anchoring effects (now measured using the

[18] Staying with the above example, if a buyer moves within the same block group, PREVIOUSPFOOT and CURRENTPFOOT are the same, which implies that ln(PFOOTRATIO_BG) is zero. It should be noted that these specifications do allow for anchoring effects to influence buyers that are moving within the same metropolitan area but across anchoring geographies. For instance, if a buyer is changing jurisdictions within the same metropolitan area, the NEWMARKET term will be equal to zero, but the ln(PFOOTRATIO_JURIS) term will likely be non-zero.

Table 4 Fixed effects regressions: price ratio specification

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
ln(MILESFROMHOME)	0.00329**	0.00313**	0.00312**			
	(0.00147)	(0.00148)	(0.00150)			
ln(PFOOTRATIO_BG)	0.0454***			0.0447***		
	(0.0107)			(0.0107)		
ln(PFOOTRATIO_JURIS)		0.0332**			0.0329**	
		(0.0134)			(0.0133)	
ln(PFOOTRATIO_COUNTY)			0.0224			0.0229
			(0.0173)			(0.0170)
ln(TOTLIV)	0.271***	0.271***	0.272***	0.271***	0.271***	0.271***
	(0.0205)	(0.0206)	(0.0206)	(0.0205)	(0.0206)	(0.0206)
ln(AGE)	−0.0184***	−0.0186***	−0.0187***	−0.0188***	−0.0189***	−0.0190***
	(0.00436)	(0.00438)	(0.00437)	(0.00437)	(0.00438)	(0.00437)
NEWMARKET				0.0203***	0.0204***	0.0205***
				(0.00659)	(0.00656)	(0.00656)
Observations	6,666	6,666	6,666	6,666	6,666	6,666
R-squared	0.814	0.813	0.813	0.814	0.813	0.813
Number of neighborhood fixed effects	2,847	2,847	2,847	2,847	2,847	2,847
TRS fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

ln(PFOOTRATIO) variables) decline as progressively larger geographies are used to define the market area.

In our final alternative specification, instead of entering the anchoring and distance measures into the regression equation separately, we generate a typology of buyers, registered by including a set of dummy variables in the equation.[19] Buyers belong to one of three categories: intra-market movers, inter-market movers relocating from a higher-cost area (NEWMARKET_HIGH), and inter-market movers relocating from a lower-cost area (NEWMARKET_LOW). A buyer is characterized as moving from a higher-cost market if PREVIOUSPFOOT is greater than CURRENTPFOOT; similarly, a buyer is said to move from a lower-cost market if PREVIOUSPFOOT is less than CURRENTPFOOT. As in the previous specifications, we define PREVIOUSPFOOT and CURRENTPFOOT over the three different geographies (also as before, we use the variable name convention of the suffix following the underscore denoting the geography).

The theory discussed above suggests that, although all new entrants into the housing market will be disadvantaged relative to intra-market movers because of their higher search costs, new market entrants from higher-priced locales are expected to be particularly disadvantaged if anchoring leads them to systematically overestimate the mean of the price distribution in their new market. In terms of our empirical model, this would suggest that estimated coefficients on NEWMARKET_HIGH and NEWMARKET_LOW should both be positive, with the NEWMARKET_HIGH coefficient greater than NEWMARKET_LOW coefficient.[20] The results (see Table 5) are in accord with these expectations: 1) regardless of the geographic definition (block group, jurisdiction, or county), inter-market movers relocating from higher-cost markets are found to pay a statistically significant premium for housing, with this premium ranging between 2 and 3% and 2) the NEWMARKET_LOW coefficient is positive and smaller than the NEWMARKET_HIGH coefficient in each of the models. Also of interest is the finding that the estimated coefficient on the NEWMARKET_LOW term is only statistically significant (and then only at the 10% level) in the model utilizing the block group anchoring geography. These results suggest that the lower price anchors of inter-market movers relocating from lower-cost locales may act to offset the search cost disadvantage they incur relative to intra-market movers.

Accounting Further for Possible Demand Effects

Our earlier comparison of the results obtained from OLS and fixed effects models suggests that the latter models control for much of the unobserved heterogeneity across buyers coming from different distances and markets. However, there remains

[19] In addition to the specifications reported here, we also estimated models in which we: (1) allowed for the NEWMARKET term to vary across markets of different size and (2) allowed for interactions between the anchoring variables and the distance variables. For the latter set of models, the interaction terms were never statistically significant; in the former set of models there did not appear to be a systematic relationship between the NEWMARKET price premium and the size of the market into which the buyer relocated. These results are available from the authors upon request.

[20] In all of the specifications reported in Table 5, intra-market movers serve as the reference category.

Table 5 Fixed effects regressions: baseline buyer category specification

Dependent variable Coefficient	ln(SALE) (1)	ln(SALE) (2)	ln(SALE) (3)
ln(TOTLIV)	0.275***	0.275***	0.275***
	(0.0206)	(0.0205)	(0.0205)
ln(AGE)	−0.0195***	−0.0195***	−0.0195***
	(0.00445)	(0.00445)	(0.00446)
NEWMARKET_HIGH_BG	0.0224***		
	(0.00821)		
NEWMARKET_LOW_BG	0.0183*		
	(0.00959)		
NEWMARKET_HIGH_JURIS		0.0245***	
		(0.00804)	
NEWMARKET_LOW_JURIS		0.0152	
		(0.00953)	
NEWMARKET_HIGH_COUNTY			0.0307***
			(0.00821)
NEWMARKET_LOW_COUNTY			0.00707
			(0.00959)
Observations	6,666	6,666	6,666
R-squared	0.809	0.809	0.809
Number of neighborhood fixed effects	2,847	2,847	2,847
TRS fixed effects?	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

the possibility that even within the small areas defined by our neighborhood fixed effects there may be buyer characteristics that are correlated with both demand and our distance and anchoring test variables. Ideally, to control for differences in buyers within neighborhoods we would include variables known to affect housing demand, such as the buyer's income, education, and age.

Although we do not have these variables at the household level, we can identify the block group from which the buyer is moving. We are subsequently able to construct proxy variables for the missing household descriptors using block group data reported in the 2000 Census. The three variables that we use to control for buyer heterogeneity are the median family income (FAMILYINC), the percentage of adults with a college degree (PERCOLLEGE), and the average age of homeowners (HOMEOWNERAGE) in the census block group from which the buyer moved. We discuss the results from adding the proxy variables to our baseline and alternative models in turn.

Table 6 reports the results from adding the proxy variables to our baseline set of models. Of the three proxy variables, the estimated coefficients on ln(FAMILYINC) and ln(HOMEOWNERAGE) are statistically significant in all models. The estimated coefficient on the proxy for educational attainment (ln(PERCOLLEGE)), however, is not statistically significant in any of the specifications. To avoid repetition, we note that these findings hold true for both the baseline and alternative models.

Focusing first on the search cost variables, a comparison of the results in Table 6 with those in Table 2 suggests that the estimated influence of these variables is robust to the inclusion of the proxy variables: the estimated coefficients on NEWMARKET are essentially the same, while the estimated ln(MILESFROMHOME) coefficients are slightly larger in Table 6 than those reported in Table 2. These results lend further credence to our conclusion that non-local buyers pay more for housing.

The anchoring estimates are more sensitive to the inclusion of the demand proxies, as the magnitudes and levels of statistical significance of the PREVIOUSPFOOT coefficients fall somewhat after the introduction of the proxies. But this result is not surprising since the anchoring variables are highly correlated with the proxy variables. However, despite the collinearity being the highest in the model using the block group anchoring variables, estimated coefficients on the ln(PREVIOUSPFOOT_BG) terms in Table 6 are still statistically significant regardless of the distance measure employed. Hence, while the inclusion of the buyer proxies weakens our results in support of the anchoring hypothesis, overall the evidence remains supportive.

Tables 7 and 8 report the results obtained for the alternative specifications with the inclusion of the demand variables. For the models in which the ln(PFOOTRATIO) variables serve as our anchoring variables (Table 7), the introduction of the block group proxy variables attenuates all of the anchoring coefficients. As in Table 6, however, the anchoring variables remain statistically significant when utilizing the block group anchoring geography and the coefficients on the distance measures exhibit little change. The models utilizing the discrete categorization of buyers (Table 8) also still provide support for our hypotheses, even with the inclusion of the proxy variables. Moving between Tables 5 and 8, we see that the inclusion of the additional controls reduces the magnitude of the NEWMARKET_HIGH variables, but increases the

Table 6 Fixed effects regressions: baseline specification with proxy variables

Dependent variable Coefficient	ln(SALE) (1)	ln(SALE) (2)	ln(SALE) (3)	ln(SALE) (4)	ln(SALE) (5)	ln(SALE) (6)
ln(MILESFROMHOME)	0.00407***	0.00414***	0.00431***			
	(0.00149)	(0.00151)	(0.00152)			
ln(PREVIOUSPFOOT_BG)	0.0325***			0.0335***		
	(0.0119)			(0.0118)		
ln(PREVIOUSPFOOT_JURIS)		0.0216			0.0235*	
		(0.0141)			(0.0139)	
ln(PREVIOUSPFOOT_COUNTY)			0.00612			0.0105
			(0.0173)			(0.0169)
ln(TOTLIV)	0.279***	0.277***	0.278***	0.277***	0.276***	0.276***
	(0.0202)	(0.0202)	(0.0203)	(0.0202)	(0.0203)	(0.0203)
ln(AGE)	−0.0186***	−0.0186***	−0.0186***	−0.0191***	−0.0191***	−0.0191***
	(0.00443)	(0.00443)	(0.00442)	(0.00443)	(0.00443)	(0.00442)
NEWMARKET				0.0188***	0.0189***	0.0194***
				(0.00663)	(0.00661)	(0.00661)
ln(FAMILYINC)	0.0471***	0.0518***	0.0526***	0.0457***	0.0504***	0.0511***
	(0.0103)	(0.0103)	(0.0103)	(0.0103)	(0.0103)	(0.0103)
ln(PERCOLLEGE)	−0.00297	−0.00129	−0.000603	−0.00399	−0.00232	−0.00167
	(0.00603)	(0.00594)	(0.00591)	(0.00601)	(0.00592)	(0.00590)
ln(HOMEOWNERAGE)	0.0526**	0.0580***	0.0596***	0.0494**	0.0549**	0.0569***
	(0.0220)	(0.0219)	(0.0220)	(0.0220)	(0.0219)	(0.0220)
Observations	6,666	6,666	6,666	6,666	6,666	6,666
R-squared	0.812	0.811	0.811	0.812	0.811	0.811
Number of neighborhood fixed effects	2,847	2,847	2,847	2,847	2,847	2,847

Table 6 (continued)

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
TRS fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 7 Fixed effects regressions: price ratio specification with proxy variables

Dependent variable Coefficient	ln(SALE) (1)	ln(SALE) (2)	ln(SALE) (3)	ln(SALE) (4)	ln(SALE) (5)	ln(SALE) (6)
ln(MILESFROMHOME)	0.00415***	0.00421***	0.00428***			
	(0.00149)	(0.00150)	(0.00152)			
ln(PFOOTRATIO_BG)	0.0250**			0.0259**		
	(0.0114)			(0.0114)		
ln(PFOOTRATIO_JURIS)		0.0168			0.0184	
		(0.0136)			(0.0134)	
ln(PFOOTRATIO_COUNTY)			0.00791			0.0122
			(0.0174)			(0.0170)
ln(TOTLIV)	0.277***	0.277***	0.278***	0.276***	0.276***	0.276***
	(0.0202)	(0.0203)	(0.0203)	(0.0203)	(0.0203)	(0.0203)
ln(AGE)	−0.0185***	−0.0185***	−0.0186***	−0.0190***	−0.0190***	−0.0191***
	(0.00442)	(0.00443)	(0.00442)	(0.00443)	(0.00443)	(0.00442)
NEWMARKET				0.0190***	0.0191***	0.0193***
				(0.00663)	(0.00660)	(0.00661)
ln(FAMILYINC)	0.0486***	0.0521***	0.0525***	0.0472***	0.0507***	0.0510***
	(0.0103)	(0.0103)	(0.0103)	(0.0103)	(0.0103)	(0.0103)
ln(PERCOLLEGE)	−0.00235	−0.00111	−0.000635	−0.00338	−0.00215	−0.00169
	(0.00602)	(0.00594)	(0.00591)	(0.00600)	(0.00593)	(0.00590)
ln(HOMEOWNERAGE)	0.0543**	0.0585***	0.0598***	0.0511**	0.0554**	0.0570***
	(0.0220)	(0.0219)	(0.0220)	(0.0220)	(0.0219)	(0.0220)
Observations	6,666	6,666	6,666	6,666	6,666	6,666
R-squared	0.811	0.811	0.811	0.811	0.811	0.811
Number of neighborhood fixed effects	2,847	2,847	2,847	2,847	2,847	2,847

Table 7 (continued)

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)	(4)	(5)	(6)
TRS fixed effects?	Yes	Yes	Yes	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 8 Fixed effects regressions: discrete buyer category specification with proxy variables

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)
ln(TOTLIV)	0.276***	0.276***	0.276***
	(0.0203)	(0.0203)	(0.0203)
ln(AGE)	−0.0190***	−0.0190***	−0.0191***
	(0.00442)	(0.00443)	(0.00443)
NEWMARKET_HIGH_BG	0.0172**		
	(0.00823)		
NEWMARKET_LOW_BG	0.0229**		
	(0.00964)		
NEWMARKET_HIGH_JURIS		0.0198**	
		(0.00812)	
NEWMARKET_LOW_JURIS		0.0200**	
		(0.00961)	
NEWMARKET_HIGH_COUNTY			0.0269***
			(0.00829)
NEWMARKET_LOW_COUNTY			0.0107
			(0.00959)
ln(FAMILYINC)	0.0517***	0.0515***	0.0503***
	(0.0103)	(0.0103)	(0.0103)
ln(PERCOLLEGE)	−0.00136	−0.00154	−0.00146
	(0.00591)	(0.00590)	(0.00590)
ln(HOMEOWNERAGE)	0.0562**	0.0561**	0.0569***
	(0.0219)	(0.0219)	(0.0220)
Observations	6,666	6,666	6,666

Table 8 (continued)

Dependent variable	ln(SALE)	ln(SALE)	ln(SALE)
Coefficient	(1)	(2)	(3)
R-squared	0.811	0.811	0.811
Number of neighborhood fixed effects	2,847	2,847	2,847
TRS fixed effects?	Yes	Yes	Yes
County-specific trend?	Yes	Yes	Yes
County-MVALUE interaction?	Yes	Yes	Yes

Standard errors clustered at the TRS level

Standard errors in parentheses

*** $p<0.01$, ** $p<0.05$, * $p<0.1$

magnitude of the NEWMARKET_LOW variables. All of the NEWMARKET_HIGH variables are statistically significant in Table 8 regardless of the anchoring geography employed. In addition, the NEWMARKET_LOW variables are statistically significant for the block group and jurisdiction geographies. These results further support our conclusion that inter-market movers pay more for housing than local movers. However, except for the county level geography, the addition of the proxy variables results in there being little difference between the estimated coefficients on the NEWMARKET_HIGH and NEWMARKET_LOW variables. Hence, in contrast to all earlier results, those reported in Table 8 provide less support for the anchoring hypothesis.[21]

Conclusion

A number of authors have suggested that non-local buyers of single-family homes incur higher search costs and have less nuanced knowledge of the local housing market. If these buyers are being accurately characterized, housing market search theory implies that they will pay more for housing. We have labeled this the "distant buyer" hypothesis. It has also been suggested that buyers coming from higher priced markets may have inflated estimates of the price distribution. Again, search theory implies that if this is true, then these buyers will pay more for housing. This has been labeled the "anchoring hypothesis." Existing evidence on both of these hypotheses is sparse and, in the case of the distant buyer hypothesis, contradictory. The issue is therefore unresolved whether existing housing market institutions provide sufficient information to eliminate systematic differences in the price of housing paid by buyers of different types.

In this paper we utilize a new data set from the state of Florida to reexamine the distant buyer and anchoring hypotheses. Our empirical models are estimated using much larger samples than those employed in previous studies and are specified to minimize possible bias resulting from omitted variables. The results strongly support the distant buyer hypothesis and therefore are contrary to the results obtained by the two previous studies that have focused on owner-occupied housing. The results also lend support to the anchoring hypothesis. However, they do not rule out the possibility that buyers coming from higher priced markets pay more because they obtain greater unobserved housing quality.

There are a number of directions for future research. Most importantly, future work should determine whether the price premia paid by distant buyers (whether from higher search costs or anchoring effects) varies with household income levels. On the one hand, because search costs are largely time costs, it is expected that

[21] At the suggestion of one of the referees, we conducted two additional robustness tests. In the first of these tests, we reestimated the equations reported in Tables 6, 7 and 8 including, in addition to proxy variables for the buyer's previous neighborhood, the same characteristics of the buyer's current neighborhood. The second robustness test involved implementing the identification strategy outlined in Harding et al. (2003), in which the characteristics of the buyer's previous and current neighborhood enter the regression equation in sums and differences. The results of these robustness tests, available upon request, produced coefficients on the anchoring and distance variables that were very similar in magnitude and significance to those reported in Tables 6, 7 and 8.

higher income households may search less and pay more. On the other hand, higher income households may more frequently acquire information prior to visiting the new market that enables them to conduct a more efficient search upon arrival. In addition, higher income households may use different anchors than lower income households. For example, the bid prices of higher income households may be less affected by priors in their old market if the households have greater access to the abundance of real estate market information available on the Internet.

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