Listing Contract Length and Time on Market

Authors

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Abstract

Miceli (1989), in a search for the optimal time to allow a broker to market property, posits that the principal (seller) may use the length of the listing contract to motivate the agent (listing broker) to better align incentives. Expanding slightly on Miceli, this work predicts that longer time allotted the broker to market residential property will decrease broker effort, resulting in lower search intensity and eventually a longer marketing span for property, ceteris paribus. This prediction is borne out across three empirical modeling methodologies commonly used in time-on-market studies.

The unobservable nature of marketing property may allow the broker to shirk, leading to a misalignment of incentives between an agent and his principal. This leads to the obvious question of how the time contractually allotted a broker to market property impacts the average selling time of these properties.[1] Building from the initial theoretical work of Miceli (1989), this research makes an original empirical investigation into that question.

Miceli (1989) models the relationship between listing contract length, the tendency for unobserved brokers to shirk, and the cost of finding another agent in the event of marketing failure, along the way to developing an optimal listing contract length. He develops most of the theoretical relationship between broker effort, broker search intensity, arrival rates, and time on market needed for this present work.[2] However, a direct theoretical prediction between listing contract length and property duration is not provided, most likely as the explanation of property duration is not the main focus of his work.

This research expands slightly on Miceli (1989) capitalizing on both: (a) the inverse relationship between arrival rates and property duration and (b) recently available data on the time allotted a broker to market property to investigate the relationship between listing contract length, hereafter often referred to as length of contract or LOC, and residential property marketing time. In particular, extending LOC leads to a decrease in broker effort, resulting in lower search intensity on the broker's part. This, in turn, leads to a lower probability of a sale and, as duration estimates are the inverse of arrival rates, to an increase in property marketing spans. In short, longer LOC leads to a longer time on market (TOM) for properties, ceteris paribus. Empirical investigation employing a variety of empirical methodologies verifies this prediction.

The finding that extended listing contracts lead to longer marketing times provides sellers of property a tool to better align their goals and those of the brokers they hire. The next section discusses the relevant literature. Sections on theory, methodology, data, empirical results, and concluding remarks follow.

Literature

The property marketing duration literature covers a wide array of topics including brokerage commissions (Zorn and Larsen, 1986), brokerage firm size (Yang and Yavas, 1995), sales price (Yavas and Yang, 1995), list price changes (Knight, 2002), and relative property size (Turnbull, Dombrow, and Sirmans, 2006). Within this body of literature, there is some agreement on the issue of joint determination of property price and selling time stemming from original works by Belkin, Hempel, and McLeavey (1976) and Miller (1978). However, the functional relationship proffered and the empirical findings in this line of the literature are quite often contradictory to one another. As an example of these contradictions, Anglin, Rutherford, and Springer (2003), using a sample of withdrawn and sold properties, argue and find support for a positive relationship between list price and property marketing time, while Turnbull and Dombrow (2007), using only sold properties, find support for an inverse relationship.

Additionally, there seems to be an ongoing disagreement over the preferred modeling technique with some favoring direct OLS modeling. Munneke and Yavas (2001), Rutherford, Springer, and Yavas (2001), and Allen and Dare (2004) provide a sample of these works. A second group seems to prefer nonlinear modeling techniques, with the Weibull distribution serving as the favorite underlying distributional assumption within this group. Anglin, Rutherford, and Springer (2003), Rutherford, Springer, and Yavas (2005), and Huang and Rutherford (2007) serve as an excellent sample of this group. A third group of authors appear to favor two-stage least squares (2SLS) estimations in order to better capture the joint determination of property price and selling time. Excellent samples here include, but are not limited to, Knight (2002), Turnbull, Dombrow, and Sirmans (2006), and Turnbull and Dombrow (2007).

To date, only three papers incorporate the time allotted a broker to market property into their analysis. Miceli (1989), as mentioned above and discussed in further detail below, is a purely theoretical piece seeking to outline an optimal contract length. Asabere, Huffman, and Johnson (1996) test the effect of listing contract length on sales price. The authors hypothesize that sellers' opportunity costs increase as time passes, leading to a reduction in reservation price. Their results suggest that transaction price falls by .04% for each day that the listing nears expiration. Clauretie and Daneshvary (2008) model principal-agent issues that arise at the end of the listing contract period. They hypothesize that as a listing contract nears expiration, brokers have an incentive to both work harder and simultaneously persuade the seller to reduce their reservation price. Their empirical analysis finds that transaction price falls as the listing contract nears expiration, supporting the hypothesis that brokers encourage their sellers to accept lower transaction prices as the listing contract nears expiration.

Building on Miceli's (1989) theoretical principal-agent model and the empirical listing contract work of Asabere, Huffman, and Johnson (1996) and Clauretie and Daneshvary (2008), this paper extends the TOM literature by testing the hypothesis that longer listing contracts result in longer TOM. Tests employed in this paper follow the empirical TOM literature cited above.

Theory[3]

Miceli (1989) investigates listing contract length as a mechanism to better align brokers' and sellers' incentives. The work develops a partial equilibrium model that balances listing contract length, the opportunity for brokers to shirk as a result of asymmetric information, and the cost to sellers of finding another broker if the property does not sell. From this equilibrium, Miceli derives an optimal listing contract length dependent upon commissions, a seller's reservation price, a seller's cost to search for another broker, and the instantaneous discount rate observed by both sellers and brokers. As part of his work, Miceli develops a relationship between listing contract length, broker effort, and broker search intensity.

Specifically, a Poisson distribution with arrival rate s is assumed for willing buyers of any seller's property with reservation price, R. Also, any broker's search intensity (s) is increasing in broker effort. Therefore, the probability of finding a willing buyer is given by:

$$\lambda = 1 - e^{-st}, \tag{1}$$

with λ increasing in s. Furthermore, duration estimates are derived by inverting λ. Thus, decreases in search intensity on the part of the listing broker brought about by decreased effort will result in longer marketing spans, ceteris paribus.

Now combining the facts that the broker is assumed to be employed by the seller under an exclusive right to sell for duration LOC (listing contract length),[4] compensation α ($0 < \alpha < 1$) is paid to the broker *iff* the property is sold, both the seller and the broker face the common instantaneous discount rate (r), and the probability of a sale at a given instant t is given by se^{-st}, the broker's revenue function can be written as:

$$\int_0^{LOC} sR\alpha e^{-(s+r)t}dt. \tag{2}$$

However, when marketing properties, brokers face costs, $c(s)$, with $c' > 0$ and $c'' \geq 0$. Considering that these costs are only faced at any point in time if there has not been a sale, the costs are weighted by the probability of no sale, e^{-st}. Thus, the broker's cost function can be expressed as:

$$\int_0^{LOC} c(s)e^{-(s+r)t}dt. \tag{3}$$

Combining Equations (2) and (3) reveals the broker's objective profit function:

$$\pi(s,T) = \int_0^{LOC} (sR\alpha - c(s))e^{-(s+r)t}dt. \tag{4}$$

Integrating across the listing period reveals:

$$\pi(s,T) = \left[\frac{sR\alpha - c(s)}{s + r}\right][1 - e^{-(s+r)LOC}]. \tag{5}$$

Solving for the first-order conditions with respect to s and setting the marginal cost of search intensity equal to marginal benefits of greater effort (induced by increased probability of earning a commission and earning that commission sooner) determines the broker's optimal level of search intensity for a given *LOC*.

$$c'(s^*) = \alpha R + \left[\frac{LOCe^{-(s^*+r)LOC}}{(1 - e^{-(s^*+r)LOC})}\right] - (s^* + r)^{-1}][s^*R\alpha - c(s^*)]. \qquad (6)^5$$

More importantly to the work at hand, changes in the optimal level of search intensity, s^*, as induced by variation in listing contract length, is derived by way of total differentiation. Specifically:

$$\partial s^*/\partial LOC \equiv (\partial^2\pi/\partial s\partial LOC)/(-\partial^2\pi/\partial s^2). \quad (7)$$

Satisfaction of the second-order conditions dictates that the denominator is positive. Therefore, the effect of *LOC* on broker search intensity is dependent upon the sign of the numerator, which is always negative for $LOC > 0$.[6] So, as listing contract length increases, the optimal level of search intensity on the broker's part declines. More specifically:

$$\frac{\partial s}{\partial LOC} < 0. \quad (8)$$

Finally, returning to the search process (Equation 1), lower effort and corresponding lower search intensity induced by providing the broker with an extended time to market property results in lower probabilities of a sale, as λ is increasing in s. And, since the hazard function (λ) is inversely related to duration estimates, longer marketing spans should be observed. More succinctly, a longer listing period negotiated between the broker and the seller results in a longer marketing time due to less intensity of search, which is the central hypothesis of this paper.

Empirical Approach

Each of the property duration modeling techniques presented in the literature review has advantages and disadvantages. OLS estimation allows for the ability to test and make corrections for self-selection and is equally flexible in dealing with endogeneity issues.[7] However, non-normal error terms can lead to biased OLS coefficient estimates, potentially rendering economic interpretation meaningless. Hazard models, often assuming a Weibull distribution of property marketing time, offer highly flexible functional specifications and are an intuitively appealing technique. Unfortunately, the aforementioned self-selection and endogeneity problems, common features in TOM studies, are not so easily dealt with in these models. Instrumental variables models, usually in the form of 2SLS, allow for the joint estimation of simultaneously determined property price and property selling time. With this technique, however, there are criticisms related to the non-normality of the error term and the difficulty in calculating required inverse Mills ratios (IMR) to control for self-selection issues between variables of interest and property marketing time.

This present study does not seek to settle these empirical modeling issues. Instead, all three methodologies are employed to maximize the robustness of any findings. In particular, this study seeks to determine whether the central hypothesis (a longer listing contract results in longer property marketing time) is consistent when tested across all commonly applied econometric techniques. To that end, the following general modeling framework is followed:

$$TOM_i = \Psi(X_i, L_i, Z_i, LOC_i, SP_i). \quad (9)$$

$$\hat{SP}_i = \delta(x_i, l_i, z_i, LOC_i). \quad (10)$$

Where TOM_i and SP_i are vectors for property marketing time and property selling price; X_i and x_i are vectors of property characteristics; L_i and l_i are vectors for location control; Z_i and z_i are vectors that include variables such as the degree of over pricing and market condition; and LOC_i, the variable of interest, is the length of time provided by the seller to the broker to market the property.

Only Equation (9) is needed for the OLS and Weibull estimations. In the 2SLS specification, SP_i is estimated in reduced form in Equation (10), with predicted values, $\hat{SP}_i$, substituting for SP_i in Equation (9). In a 2SLS setting, (9) and (10) form the system of equations between property price and property marketing time. Variables included in the TOM equation (vectors X_i, L_i, and Z_i) follow existing literature and are discussed in the Empirical Results section. In the interest of brevity, reduced form sales price results are not reported. As a reduced form equation, variables in Equation (10) (vectors x_i, l_i, and z_i) include all vectors from Equation (9) plus instruments for price.

Discussion of some of the specified regressors and other modeling issues is warranted at this point. In order to properly specify the degree of overpricing (*DOP*) as suggested in Anglin, Rutherford, and Springer (2003), a list price equation is estimated. *DOP* is the residual from that estimation and represents the percentage deviation of a property's list price from the expected list price for a typical property in the sample with the same characteristics.[8] Also, when TOM analysis is restricted to properties that have been sold, the duration data are ''censored'' (Anglin, Rutherford, and Springer, 2003). Therefore, the omission of properties that are withdrawn by the seller or for which the contract expires without a sale may result in an artificially reduced TOM in the data sample. Accordingly, this study incorporates unsuccessfully marketed properties in the hazard model (Weibull) analysis. Thus, four estimates (OLS, Weibull, Weibull with Censoring, and 2SLS) of property duration are ultimately exhibited. Finally, in order to control for market conditions, the average 30-year fixed rate mortgage rate at the sales date is specified as a proxy for market condition. Market rates are determined in part by demand conditions for property and intuitively seem a sound proxy for overall market conditions.

An additional issue to consider is the relationship between contract length and TOM.[9] Ex post, TOM cannot cause contract length simply because it occurs after the contract is negotiated. Ex ante, however, expected TOM may influence optimal contract length from both the broker's and seller's point of view. Note that explanatory variables in a typical TOM equation are specified to control for factors (e.g., property characteristics, location, market conditions, and degree of overpricing) that would determine marketing time expectations of sellers and brokers. Therefore, the coefficient of LOC is an estimate of the partial effect of listing contract length on marketing time, holding determinants of expected TOM constant.

Additional details about the operational model forms of the three techniques are omitted here as they can be determined from the reported empirical results discussed below.

Data

The data set for this paper is cross sectional and comes from a central Virginia multiple listing service (MLS). Transactions occurred over eight quarters from mid-2006 through mid-2008. After culling incomplete or miscoded data, the final sample consists of 3,502 properties, 2,186 of which were sold and 1,316 of which were unsuccessful in their marketing attempt.[10]

The MLS data includes the date at which the contract was last extended, but no information about whether the contract had been extended one or more times previous to that date. For those properties, then, it is impossible to definitively determine the original contract length and the number or length of additional extensions. Therefore all properties for which the listing contract had been extended were removed from the data set. An additional reason to exclude properties with extended contracts was discussed above. When a contract is extended, the seller and broker may use information from the property's current marketing effort when negotiating the length of the extension. That is, TOM can cause LOC. Therefore, removing all observations for which the current contract has been extended removes an element of potential endogeneity from the model.

Exhibit 1 cumulatively summarizes data for both *SOLD* (closed transactions) and *UNSOLD* (unsuccessfully marketed) properties. Variables are briefly explained in Exhibit 2. *SOLD* properties had a significantly lower list price ($180,000 compared to $231,000), were smaller, were on smaller lots, and had fewer bedrooms and baths than properties that were *UNSOLD*. Additionally, *UNSOLD* properties, not surprisingly, were overpriced relative to the sample (*DOP* = .04). Finally, properties that sold exhibited slightly longer listing contracts (166 days vs. 159).

The local real estate market slowed during the 2006–2008 sample period. There were more properties on the market in the first half of the period (2,247 in the first four quarters vs. 1,255 in the last four). Additionally, more properties sold in the first half (68.5% of the sales occurred in the first four quarters and 31.5% in the second), while the proportion of unsold properties were more evenly distributed (57% of the *UNSOLD* properties occurred in the first four quarters to

Exhibit 1 | Descriptive Statistics

	SOLD Properties		*UNSOLD* Properties		
Variable	Mean	Std. Dev.	Mean	Std. Dev.	*t*-Stat.
TOM	81.18	48.38	137.01	75.012	−26.763
LOC	166.13	77.83	159.29	80.02	2.492
Selling Price	176249	93579			
DOP	−0.020	0.219	0.040	0.274	−7.158
Square Feet	1848.62	762.76	2087.48	916.740	−8.309
Age	27.548	27.482	26.737	33.946	0.773
Acreage	0.647	0.609	0.778	0.719	−5.785
Bedrooms	3.141	0.770	3.284	0.868	−5.068
Full Bathrooms	1.947	0.703	2.093	0.767	−5.788
Half Bathrooms	0.412	0.527	0.464	0.537	−2.801
Fire	0.623	0.485	0.660	0.474	−2.213
Hardwood	0.536	0.499	0.559	0.497	−1.302
Full basement	0.566	0.496	0.538	0.499	1.620
Vacant	0.340	0.474	0.305	0.461	2.095
30-Year FRM-SD	6.285	0.240			
30-Year FRM-LD	6.332	0.241	6.335	0.246	−0.370
Listing Price	179686	95846	231175	158079	−12.001

Notes: Means, standard deviations, and *t*-tests for differences between the means for *SOLD* (n = 2,186) and *UNSOLD* (n = 1,316) Properties.

43% in the second). Said another way, fewer properties were on the market and a smaller percentage of them sold as the market slowed.

Exhibit 3 provides a frequency distribution of contract lengths and mean market duration. Not surprisingly, *UNSOLD* properties were on the market longer across all contract length categories. This result is also illustrated in average *TOM* between *SOLD* and *UNSOLD* properties in Exhibit 1. As shown in Exhibit 3, a plurality of listing contracts (1,077) exhibited durations between three and six months, with an average *TOM* of just under 70 days. After consultation with several brokers in the area covered by the MLS, there was agreement that these contract lengths are representative for the time period examined.

Finally, the listing contract and TOM data used in this study represent the most recent attempted sale of the home and do not reflect the full marketing history of a property.[11] However, by excluding properties that sold within a few days and properties for which the current contract had been extended, and by incorporating

Exhibit 2 | Variable Legend

Variable Name	Variable Definition
TOM	Property time-on-market, days from list date to under contract date.
LOC	Length of listing contract provided the broker to market property.
Selling Price	Property transaction price.
DOP	Degree of Overpricing as defined in Anglin, Rutherford, and Springer (2003).
Square Feet	Property size in square feet.
Age	Property age in years.
Acreage	Property acreage.
Bedrooms	Number of bedrooms.
Full Bathrooms	Number of full bathrooms.
Half Bathrooms	Number of half bathrooms.
Fire	1 if the property has a fireplace, 0 otherwise.
Hardwood	1 if the property has hardwood floors, 0 otherwise.
Full basement	1 if the property has a full basement, 0 otherwise.
Vacant	1 if the property is vacant, 0 otherwise.
30-Year FRM-LD	30-year fixed rate mortgage at listing date.
30-Year FRM-SD	30-year fixed rate mortgage at under contract date.
Listing Price	Listing price of the property.
SOLD	1 if the property sold, 0 otherwise.

expired and withdrawn properties into the analysis, this research employs data that minimize potential biases due to incomplete marketing histories.[12]

Empirical Results

The results shown in Exhibit 4 provide support for this work's central hypothesis that a longer listing contract period, on average, results in a longer *TOM*. The results are robust across each of the four analyses and demonstrate that *LOC* is a positive and significant determining factor in *TOM*.

Other results revealed in Exhibit 4 are generally consistent with prior studies. The simultaneous determination of price and *TOM* is found to be inversely related, as suggested by Turnbull and Dombrow (2007). The degree of overpricing (*DOP*), which is generally a proxy for seller motivation, is positively related to *TOM* in all results, and significant in all but the 2SLS model, illustrating that overpriced properties require longer marketing periods. This result is consistent with Anglin, Rutherford, and Springer (2003). Variables that capture housing characteristics have expected signs and are generally significant in the OLS[13] and Weibull

Exhibit 3 | Frequency Distribution of *LOC* & *TOM*

Contract Length	*SOLD* Mean *TOM*	Number of Properties	*UNSOLD* Mean *TOM*	Number of Properties
LOC < 91 days	50.83	185	64.69	228
90 < *LOC* < 181	69.55	1,077	115.79	516
180 < *LOC* < 271	89.66	758	163.55	462
270 < *LOC* < 366	156.50	107	308.86	71
LOC > 365	143.07	59	213.23	39

analyses. In the interest of brevity, a detailed discussion of these commonly found effects is omitted.

In both Weibull models (middle columns of Exhibit 4) the coefficient of *LOC* is positive indicating that an increase in contract length increases *TOM*, or, alternatively, decreases the hazard rate. However, direct comparison of the two Weibull models is impossible because of required changes in specification (omission of sales price and replacement of the mortgage rate at sales date with mortgage rate at list date) for the sample that includes unsold properties. A speculative observation, however, is that significance levels of the overall equation and of two key explanatory variables (*LOC* and *DOP*) are higher for the model that incorporates the impact of censoring. This is suggestive of support for the conclusion by previous authors, such as Anglin, Rutherford, and Springer (2003), that modeling the impact of censoring is a superior methodology.

The uniqueness of the 30-year mortgage rate (market condition proxy) and *SOLD* and *UNSOLD* properties deserves some discussion in order to clarify the measurement of this metric and understand its role in residential property marketing duration. The 30-year mortgage rate displays a negative and significant effect in the OLS, Weibull, and 2SLS estimations; however, only a 10% significance level is produced in the latter. Rising rates in active markets are generally associated with markets in which homes sell relatively quickly, making 30-Year FRM-SD a valid proxy control for market condition. Note that in the full sample (censoring) Weibull model, the 30-year mortgage rate at the list date has a negative but not significant coefficient. It might well be the case that market rates are less likely to affect properties that did not successfully market. Perhaps, these properties were simply overpriced. The difference between the *DOP* for *SOLD* and *UNSOLD* properties in Exhibit 1 provides support for this possibility.

Returning to the variable of interest, *LOC* significantly increases *TOM*, in the OLS and 2SLS estimates (Exhibit 4), revealing about a 5% increase in *TOM* for each 10% increase in *LOC*.[14] Similarly, the estimated Weibull coefficients for *LOC* are positive and significant at the 1% level. The Weibull coefficients, 0.15 and 0.17,

Exhibit 4 | OLS, Weibull, and 2SLS Estimations of Property Duration

Dependent Variable: *LnTOM*

	OLS	*t*-value	Weibull	*t*-value	Weibull–w/Censoring	*t*-value	2SLS	*t*-value
Constant	8.40	5.35***	1.90	6.20***	0.46	6.52***	45.97	1.27
LN(LOC)	0.52	19.05***	0.14	28.73***	0.17	49.73***	0.53	16.22***
LN(Selling Price)	−0.81	−3.74***	−0.13	−3.22***			−6.34	−1.19
DOP	0.85	3.73***	0.14	3.26***	0.02	3.94***	6.38	1.20
LN(Square Feet)	0.71	4.35***	0.12	3.90***	0.04	5.85***	4.64	1.37
LN(Age)	−0.18	−5.48***	−0.03	−4.45***	−0.006	−4.73***	−0.98	−1.42
LN(Acreage)	0.01	1.46	0.001	0.69	0.0002	0.18	0.11	1.33
Bedrooms	−0.04	−1.98**	−0.009	−2.58***	−0.002	−0.95	−0.19	−1.30
Full Bathrooms	0.07	2.38**	0.011	2.04**	−0.006	−2.03**	0.43	1.24
Half Bathrooms	0.04	1.62	0.007	1.54	−0.0002	−0.07	0.33	1.19
Fire	0.11	2.63***	0.018	2.34**	−0.005	−1.43	0.90	1.18
Hardwood	0.10	3.46***	0.018	3.19***	0.003	1.24	0.59	1.26
Full basement	0.03	1.17	−0.002	−0.45	−0.011	−3.92***	0.29	1.16
Vacant	0.00	0.17	0.004	1.02	0.001	0.24	−0.28	−1.02

Exhibit 4 | (continued)

OLS, Weibull, and 2SLS Estimations of Property Duration

Dependent Variable: *LnTOM*

	OLS	*t*-value	Weibull	*t*-value	Weibull–w/Censoring	*t*-value	2SLS	*t*-value
30-Year FRM-SD	−0.28	−4.40***	−0.05	−4.76***			−0.19	−1.77*
30-Year FRM-LD					−0.0003	−0.04		
N	2,186		2,186		3,502		2,186	
F	26.06							
R^2	.2662						.2560	
Chi Sq.			1,688.61		3,965.82		647.99	

Notes: Time trend and location controls have been suppressed for the purpose of exposition. *SP* is instrumented and specified as $Ln\hat{SP}$ (not *LnSP*) in the 2SLS model. All models use Huber-White robust standard errors in the estimation of *t*-statistics. Both Weibull models employ the accelerated failure time metric.

*Significant at the .10 level.

**Significant at the .05 level.

***Significant at the .01 level.

Exhibit 5 | OLS, Weibull, and 2SLS Estimations of Property Duration

Dependent Variable: *LnTOM*

Categories of *LOC*

	OLS	*t*-value	Weibull	*t*-value	Weibull–w/Censoring	*t*-value	2SLS	*t*-value
Constant	11.62	7.00***	2.94	8.73***	1.39	17.01***	42.74	1.15
LOC < 91 Days	−0.74	−8.69***	−0.19	−11.97***	−0.21	−18.33***	−0.76	−8.54***
LOC 91 to 180 Days	−0.45	−5.46***	−0.12	−7.40***	−0.12	−11.37***	−0.46	−5.40***
LOC 181 to 270 Days	−0.26	−3.17***	−0.07	−4.27***	−0.06	−5.69***	−0.25	−3.07***
LOC 271 to 365 Days	0.14	1.34	0.03	1.60	0.04	3.73***	0.15	1.48
LN(Selling Price)	−0.81	−3.51***	−0.15	−3.23***			−5.38	−1.15
DOP	0.83	3.46***	0.15	3.18***	0.01	2.03**	5.40	1.15
LN(Square Feet)	0.72	4.17***	0.14	3.92***	0.04	5.21***	3.96	1.19
LN(Age)	−0.19	−5.27***	−0.03	−4.70***	−0.007	−4.72***	−0.85	−1.25
LN(Acreage)	0.02	1.50	0.001	0.76	0.0005	0.40	0.10	1.16
Bedrooms	−0.04	−2.08**	−0.009	−2.45**	−0.003	−1.12	−0.17	−1.29
Full Bathrooms	0.07	2.19**	0.011	1.91*	−0.006	−1.71*	0.36	1.19
Half bathrooms	0.04	1.56	0.006	1.13	−0.002	−0.56	0.28	1.14
Fire	0.11	2.53**	0.017	2.03**	−0.009	−2.15**	0.76	1.13
Hardwood	0.10	3.15***	0.018	2.78***	0.001	0.36	0.50	1.21
Full basement	0.02	0.80	−0.002	−0.34	−0.013	−3.61***	0.23	1.06
Vacant	0.02	0.70	0.010	2.06	0.005	1.46	−0.22	−0.89

Exhibit 5 | (continued)

OLS, Weibull, and 2SLS Estimations of Property Duration

Dependent Variable: *LnTOM*

Categories of *LOC*

	OLS	*t*-value	Weibull	*t*-value	Weibull–w/Censoring	*t*-value	2SLS	*t*-value
30-Year FRM-SD	−0.34	−4.96***	−0.07	−6.00***			−0.26	−2.73***
30-Year FRM-LD					−0.003	−0.32		
N	2,186		2,186		3,502		2,186	
F	22.70							
R^2	.2662							
Chi Sq.			1,284.63		2,833.49		574.3	

Notes: Time trend and location controls have been suppressed for the purpose of exposition. *SP* is instrumented and specified as $Ln\hat{SP}$ (not *LnSP*) in the 2SLS model. All models use Huber-White robust standard errors in the estimation of *t*-statistics. Both Weibull models employ the accelerated failure time metric.

*Significant at the .10 level.

**Significant at the .05 level.

***Significant at the .01 level.

are much smaller than those estimated in the OLS and 2SLS regressions, suggesting that those models may overstate the impact of longer contracts. Differences in magnitude between estimated coefficients of Weibull models and OLS or 2SLS have been noted by other authors including Allen, Faircloth, and Rutherford (2005). However, the results are robust across commonly-used methodologies and provide a preponderance of statistical evidence in support of the central hypothesis that increasing the time allotted a broker to market property results in increased property selling time.

Additional detail on the impact of varying contract lengths on *TOM* is provided by creating categorical variables for *LOC* (results are reported in Exhibit 5). Categorical variables were created for contracts less than 91 days, 91 to 180 days, 181 to 270 days, 271 to 365 days, and greater than 365 days (the omitted category in all results). OLS estimates illustrate the common result found across all four techniques. Properties with the shortest contracts ($<$ 91 days) had the shortest *TOM*, ceteris paribus, relative to properties with contracts greater than one year. *TOM* increases as *LOC* categories increase up to 270 days. The 271 to 365 day category shows longer *TOM* than the $>$ 365-day category, but the difference in coefficients of the two longest contract categories is not statistically significant. These results are suggestive that listing contract length exhibits a declining marginal effect on property marketing spans. These empirical results extend beyond Miceli's (1989) model and may warrant further research.

These results as a whole represent an empirical confirmation of Miceli's (1989) model and are therefore entirely consistent with the hypothesis that length of the listing contract can be an effective tool to align broker and seller incentives.

Conclusion

This paper contributes to the time-on-market literature by empirically investigating, for the first time, the relationship between length of listing contract and time-on-market. While the theoretical foundation for this study was put in place almost two decades ago, the recent availability of data concerning the time allotted a broker to market a property now makes this study possible. This work hypothesizes that extending the listing contract period results in a decrease in broker effort, resulting in lower search intensity on the broker's part and corresponding lower arrival rates, and, as duration estimates are the inverse of arrival rates, an increase in marketing time. Empirical investigation of this central hypothesis across a variety of methodologies verifies this relationship.

These results contribute to the rich literature on principal-agent incentive misalignments in real estate brokerage by examining the impact of listing contract length. Empirical research on listing contracts, which began with Asabere, Huffman, and Johnson (1996) and extended by Clauretie and Daneshvary (2008), focus on pricing issues as contracts expire, while this study concentrates on overall marketing time.

The findings help to further the understanding of the unique relationship between property sellers and their brokers, hopefully providing both sellers and brokers alike with needed information to assist them in the listing and selling process. This additional insight into the effect of listing contract length may help to better align broker and seller incentives. However, several interesting questions regarding the length of the listing contract remain unanswered, including the interrelationships between listing contract length and probability that a home will sell, the impact of dual agency, and the degree of overpricing.

Endnotes

1 In this paper, the term broker is used generically for all licensees including brokers and agents.

2 Time on market, property marketing time, selling time, marketing span, and duration are used as synonyms in this work.

3 Notation employed in this work closely, but not exclusively, follows Miceli (1989) for ease of comparison.

4 Here *LOC* is the equivalent of the termination date of the listing. In Miceli (1989), *T* is its analog. Therefore, a listing period encompasses the time period t_0 to *LOC*.

5 See Endnote 15 in Miceli (1989) for an explanation of second order conditions.

6 Again, see Equation (5) in Miceli (1989) for a more exacting explanation.

7 That is to say, Heckman and Hausman tests for self-selection and endogeneity are easily incorporated into OLS estimations of property duration.

8 Results of the list price regression are not reported here but are available from the authors upon request.

9 The authors thank anonymous referees for raising this issue.

10 The authors are indebted to an anonymous referee for pointing out that properties which are on the market for an extremely short time [aka: *NOMARKET* or *QUICK*—see Sirmans, Turnbull, and Dombrow (1995)] may be the result of unrepresentative marketing and inclusion of these properties may bias analysis of the length of contract—TOM relationship. Therefore, the properties with extremely short TOM are deleted from the sample.

11 For instance, an alternative explanation for a positive relationship between *LOC* and *TOM* would be that a property's unique marketing history, if known by the seller or broker, could influence current expected marketing time and, therefore, contract length. If a property was known to have sold slowly in the past, expectations of current marketing time might be influenced and a longer listing contract could be in the interest of both seller and broker. If past marketing time is specific to the property (rather than due to market conditions that existed at that time) and those unique property characteristics are not perfectly captured by property and market control variables, then some of this effect may be captured by the LOC coefficient.

12 As for past marketing histories of property, this work explicitly assumes, as all prior duration studies implicitly assume, that any error induced by the omissions of property marketing histories is independent and identically-distributed and thereby does not impact the overall distribution of property marketing time. In short, the authors of this

work recognize the inherent data issues and the implied limitation in this type of work but choose to continue property duration research as it is important to gain an increased understanding of this key market outcome.

[13] To address the potential for sample selection bias in the OLS estimate, the authors conduct a two-stage Heckman procedure. Results, not reported here but available upon request, are not materially different from the reported OLS specification.

[14] During the course of this study, numerous alternative specifications across all three estimation techniques were specified and exhibited robustness with the model results presented herein. Additionally, a previous version of this study employed data from a different MLS. That data set was comprised of only properties that sold, making it impossible to correct for selection bias in OLS regressions or censoring in Weibull. We are indebted to an anonymous referee who suggested we find a more complete data set. Results from the original data provide additional evidence of the robustness of the reported results across different times, different locales, and different empirical techniques. Specifically, LOC was a positive and statistically significant predictor of TOM in OLS, Weibull and 2SLS regressions. These estimations are available upon request from the authors.

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