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# Commercial Office Space: Testing the Implications of Real Options Models with Competitive Interactions

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We test the implications of real option pricing models with competitive interactions for commercial real estate development. The competitive nature of a local commercial real estate market relies on a Herfindahl ratio derived from individual developers' shares of total office construction in their market. All else being equal, greater competition among local developers is associated with more building starts. Other variables suggested by the real options pricing model, including the volatility of local lease rates, are also found to be statistically important. In addition, we provide evidence consistent with greater competition attenuating the extent to which increases in volatility delay commercial real estate development.

## Introduction

There are numerous applications of real option pricing models to investment decision making in commercial real estate. Examples include, among others, Titman (1985), Williams (1991, 1997), Quigg (1993), Childs, Riddiough and Triantis (1996) and Holland, Ott and Riddiough (2000).[1] However, by treating the exercise price as exogenously given, traditional real option pricing models do not take into account the fact that the exercise of the option ("investment") by one developer may affect the building price faced by other developers and so influence their exercise strategy. Recently researchers have begun to systematically investigate the effects of one developer's exercise of this option

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[1] Many other articles investigate the role of options in real estate without focusing primarily on commercial real estate development. These include, among others, Capozza and Helsley (1990), Shilling et al. (1990), Capozza and Sick (1994), Geltner, Riddiough and Stojanovic (1996), Riddiough (1997), Dale-Johnson (2001), Capozza and Li (2001, 2002), Somerville (2001) and Cauley and Pavlov (2002).

on the investment decisions of others.[2] In particular, Grenadier (2002, 2005) has developed a model to value real estate leases, which explicitly takes these competitive interactions into account.

This article takes advantage of an extensive commercial real estate database to empirically investigate the implications of real option pricing models with competitive interactions for commercial real estate development. Competition in a particular market is measured by relying on a Herfindahl concentration ratio based on individual developers' shares of total office construction in their market. We focus on a developer's option to develop additional office space and investigate the effects of various explanatory variables, including the degree of competition, on the trigger point at which such development occurs. Movements in the trigger point are proxied by the number of buildings starts, with a greater number of starts corresponding to a lower trigger point. We then statistically test whether these variables do indeed influence the observed number of building starts in the direction posited by these real option pricing models.

Our empirical results are consistent with the observed number of building starts being systematically affected by these variables and in the directions predicted by the real options theory. For example, an increase in the volatility of lease rates, all else being equal, results in a statistically reliable decrease in the number of building starts. The number of building starts is also influenced by the competitive nature of the local commercial real estate market. All else being equal, more competition amongst local developers, as measured by a correspondingly lower Herfindahl ratio, results in a greater number of building starts or, equivalently, a lower trigger point at which the option to develop is exercised. Furthermore, consistent with Grenadier's model, we find evidence that the effect of volatility on a developer's option to delay is attenuated by greater competition in a particular market.

Bulan, Mayer and Sommerville (2002) also investigate the effects of competition on option values in commercial real estate. In particular, they examine condominium development in Vancouver, Canada, between 1979 and 1998 and find that increases in risk, both systematic as well as unsystematic, delay condominium investment. They also find that an increase in competition, measured by the actual number of future developments that will be built nearby, attenuates this relation. Our articles differ in more than how the degree of competition prevailing in a particular market is measured. For example, Bulan, Mayer and Sommerville (2002) look at only one market but over a longer period of time. We, by contrast, consider a number of different markets over time which allows

[2] There is a growing literature outside of real estate on real options with competitive interactions. Examples include, among others, Williams (1993), Smit and Ankum (1993), Huisman and Kort (1999), Lambrecht and Perraudin (2003), and Miltersen and Schwartz (2004).

us to exploit these cross-sectional data in testing the implications of the real option pricing model.

The plan of this article is as follows. The next section relies on Grenadier's model to provide a framework in which to summarize the comparative statics of the trigger level at which commercial real estate investment will occur with respect to variables suggested by real option pricing models with competitive interactions. The section "Data" discusses the data and details the dependent and independent variables used to test the implications of these models. The section "Empirical Method and Results" puts forward our empirical methodology and discusses our empirical results. We conclude in the section "Conclusion."

## The Comparative Statics of a Real Option Pricing Model with Competitive Interactions

In this section we briefly overview the implications of real option pricing models with competitive interactions for commercial real estate development. We couch our discussion in the context of Grenadier's (2002, 2005) model. This model allows us to succinctly capture within a real options framework the effects of competition on a developer's decision to build. Many of these effects, however, are not unique to Grenadier's model but characterize real option pricing models with competitive interactions in general.

A local real estate market is assumed to be oligopolistic and made up of $n$ identical developers who develop and lease identical office buildings. To fix matters, at time $t$, developer $i$ owns $q_i(t)$ units of completed and rentable space. The space is infinitely divisible, and a continuous time framework is assumed. At any point in time, developers can develop new rentable units at a constant cost of $K$ per unit of space. This investment decision is irreversible. The model also abstracts from the issue of land use choice and concentrates only on determining the optimal size of the development.

The value of owning an office building arises from its underlying service flow. The instantaneous lease rate, $P(t)$, is the price of the flow of these services. It is assumed that the lease rate evolves in such a way as to clear this market at each point in time. Following Dixit and Pindyck (1994), the market inverse demand function is assumed given by

$$P(t) = X(t)Q(t)^{-\frac{1}{\gamma}}, \tag{1}$$

where the price elasticity of demand, $\gamma$, satisfies[3] $\gamma > \frac{1}{n}$ and $Q(t) = \sum_{j=1}^{n} q_j(t)$ and is the industry supply process. Here $X(t)$ represents a multiplicative demand

[3] This is necessary to ensure a well-defined equilibrium. See Grenadier (2002).

shock. Examples of demand shocks include, among others, changes in job growth, changes in industrial production and changes in disposable income. The demand shock itself evolves as a geometric Brownian motion:

$$dX(t) = \alpha X(t)\,dt + \sigma X(t)\,dZ, \tag{2}$$

where $\alpha$ is the instantaneous conditional expected percentage change in $X(t)$ and $\sigma$ is the instantaneous conditional standard deviation. The risk-free interest rate $r$ is assumed to be constant with $r > \alpha$ to ensure convergence. The cash flows are valued in a risk-neutral framework. That is, the process for $X(t)$ is assumed to be risk adjusted.

Under the above assumptions, Grenadier (2005) derives the corresponding symmetric Nash equilibrium development strategy. In particular, he obtains the equilibrium value of each identical office building in closed-form:

$$G(X, Q) = \frac{XQ^{\frac{\gamma-1}{\gamma}}}{n(r-\alpha)} + B(Q)X^{\beta} \tag{3}$$

where

$$\begin{aligned}
\beta &= \frac{-\left(\alpha - \frac{1}{2}\sigma^2\right) + \sqrt{\left(\alpha - \frac{1}{2}\sigma^2\right)^2 + 2r\sigma^2}}{\sigma^2} > 1\\
B(Q) &= \left(\frac{v_n^{-\beta}}{n}\right)\left(\frac{\gamma}{\gamma-\beta}\right)\left[K - \left(\frac{v_n}{r-\alpha}\right)\left(\frac{\gamma-1}{\gamma}\right)\right] Q^{\frac{\gamma-\beta}{\gamma}}\\
v_n &= \left(\frac{\beta}{\beta-1}\right)\left(\frac{n\gamma}{n\gamma-1}\right)(r-\alpha)K\\
X^*(Q) &= v_n Q^{\frac{1}{\gamma}}.
\end{aligned}$$

The first term in Equation (3) represents the present value of the growing perpetuity of cash flows generated by the commercial real estate assets in place. The second term, by contrast, gives the value of the option to develop additional space.

The equilibrium strategy for each developer is to develop an incremental unit whenever the state variable $X(t)$ rises to the trigger level $X^*(Q(t))$. This solution implies that the equilibrium lease rate also follows a geometric Brownian motion but with an upper reflecting barrier at $v_n$:

$$dP(t) = \alpha P(t)\,dt + \sigma P(t)\,dZ \quad \text{when } P(t) < v_n \tag{4}$$

$$dP(t) = 0 \quad \text{when } P(t) = v_n. \tag{5}$$

In other words, so long as $X$ is sufficiently low, there will be no new investment and lease rates will evolve according to Expression (4). Otherwise, the lease rate will be fixed, Expression (5), and developers will invest.

Grenadier's model is sufficiently rich in its implications for commercial real estate investment to allow us to explore the role of various economic variables on the decision to develop additional office space. In particular, the model's implications regarding the trigger level $X^*$ at which additional investment occurs can be derived in a straight forward fashion.

For example, the trigger level is a decreasing and convex function in $n$, the number of developers:

$$\begin{aligned} \frac{\partial X^*(Q)}{\partial n} &= \frac{-X^*(Q)}{n(n\gamma - 1)} < 0 \\ \frac{\partial^2 X^*(Q)}{\partial n^2} &= \frac{2\gamma X^*(Q)}{n(n\gamma - 1)^2} > 0. \end{aligned} \tag{6}$$

Increasing competition leads developers to develop sooner as the fear of pre-emption diminishes the value of their "option to wait." As a result, we would expect to observe more building starts in the face of greater competition. This conclusion, however, is also consistent with the standard microeconomic result that a monopolist ($n = 1$) sells a smaller quantity of a good than a corresponding firm in a purely competitive market ($n \to \infty$). Empirical evidence is required to discriminate between these two alternatives.

The trigger level is a decreasing function of $\alpha$, the expected rate of growth in the level of demand:

$$\frac{\partial X^*(Q)}{\partial \alpha} < 0. \tag{7}$$

In other words, when demand is growing faster, all else being equal, the developer invests sooner.

The trigger level can also be seen to be an increasing function of the volatility of the demand shock $X(t)$:

$$\frac{\partial X^*(Q)}{\partial \sigma^2} > 0. \tag{8}$$

As in all option models without competitive interactions, an increase in volatility delays the point at which an American option is exercised.

But volatility can influence the trigger level for reasons other than its effect on the value of a developer's option to wait. In particular, if volatility risk is priced in capital markets,[4] then an increase in volatility would raise the discount rate used to value projects. As a result, a risk-adjusted discounted cash flow model would also suggest fewer building starts in the face of higher volatility.

However, distinct from a risk-adjusted discounted cash flow model, a real options model implies that this volatility effect is attenuated by the presence of competitors:

$$\frac{\partial\left(\frac{\partial X^*(Q)}{\partial\sigma^2}\right)}{\partial n} < 0. \tag{9}$$

That is, as the number of competitors $n$ increases, investment delays in Grenadier's model are unambiguously reduced. By empirically verifying this particular effect, we would provide evidence consistent with the real option pricing model as opposed to the risk-adjusted discounted cash flow model.

The effect of interest rates on the trigger level is ambiguous, that is,

$$\frac{\partial X^*(Q)}{\partial r} \lesseqgtr 0. \tag{10}$$

This result is also consistent with the real options model of Capozza and Li (2001) which shows that under certain conditions an increase in interest rates can actually increase real estate investment. Of course, an increase in interest rates unambiguously reduces real estate investment under the risk-adjusted discounted cash flow model.

Finally, the trigger function is increasing in $Q$:

$$\frac{\partial X^*(Q)}{\partial Q} > 0. \tag{11}$$

As expected, all else being equal, the larger the existing supply of office space, the higher the trigger level.

### *Caveats*

While Grenadier's model is indeed rich in its implications for commercial real estate investment, it is nonetheless a highly simplified description of actual markets. For example, the model assumes that at each point in time the supply of office space equals its demand, with the lease rate equilibrating supply with demand. This implies an absence of vacancies in the model, yet in reality vacancies are observed with some degree of regularity. Actual vacancies reflect

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[4] See, for example, Ghysels, Santa-Clara and Valkanov (2005).

the time required to search for new office space and the costs incurred in moving from one location to another, factors not addressed in Grenadier's model. In addition, actual lease rates have been observed to be somewhat sticky and not adjusting quickly to changing market conditions.

This continuous time model also assumes that construction is *instantaneous* and that it can be accomplished in *infinitesimal* amounts. In reality there are construction lags, which suggests that lagged values of the explanatory variables may effect current construction. Also, real estate investment is lumpy and so we can expect the empirical fit of the model to be far from perfect. In addition, the model assumes that developers are *homogeneous* and identical in all respects including constructing identical buildings, allowing a symmetric Nash equilibrium solution. However, actual developers vary in many respects including, among others, their size, efficiency and preferences for particular building types.

Novy-Marx (2002) argues that Grenadier's conclusion that competition erodes real option values and reduces investment delays rests critically on his assumption that developers are homogeneous and can add capacity in arbitrarily small amounts without incurring adjustment costs. To make his point, Novy-Marx provides a model of an industry with a large number of competitive firms in which opportunity costs as well as variable costs are incurred when altering capacity and, as a result, real option values become significant, investment decisions are delayed and investment will be lumpy. The implication for commercial real estate investment is that even though a large number of firms may compete vigorously, underlying heterogeneity prevents them from all competing directly over any given investment opportunity.[5]

## Data

Our data source is the CoStar Office Report, which is derived from a database maintained by the CoStar Group of existing and under-construction office buildings in a number of U.S. metropolitan areas. It covers both single-tenant as well as multi-tenant buildings and includes office, office condominium, office loft and office medical buildings, encompassing approximately 22 billion square feet of space in over 800,000 properties. The data are available on a quarterly basis and are compiled separately for class A versus class B versus class C buildings.[6] We restrict our empirical analysis to class A and B buildings. Because class C buildings tend to be older buildings, we observe little new

[5] As emphasized by Novy-Marx (2002), Grenadier's model predicts that competition should affect the level of capacity rather than the rate of investment which adds to capacity. This prediction is better tested across industries rather than concentrating on a single industry as we do.

[6] According to CoStar, class A buildings are extremely desirable investment-grade properties that command the highest rents or sale prices in a market. Class B buildings, by

construction activity for class C buildings in the CoStar Office Report. For example, the median number of class C building starts in our sample is zero per quarter.

While our sample concludes with the second quarter of 2002, CoStar's coverage begins at different times in different metropolitan areas. Because of this, our sample commences with the third quarter of 1998, or later if the data is not then available. Specifically, for the following 14 metropolitan areas, data are available from at least the third quarter of 1998: Atlanta, Baltimore, Boston, Chicago, Dallas/Fort Worth, Los Angeles, Northern New Jersey, Orange County (CA), New York City, Philadelphia, Phoenix, San Francisco, Washington D.C. and Westchester County/Southern Connecticut. Data are available from at least the first quarter of 2000 but later than the third quarter of 1998 for the following 14 metropolitan areas: Charlotte, Cincinnati, Cleveland, Columbus, Denver, Detroit, Houston, Orlando, Sacramento, St. Louis, San Diego, Seattle, South Florida, and Tampa.[7] Taken together, our sample will be restricted to these 28 metropolitan areas.

### *Description of Variables*

Real option pricing models with competitive interactions have implications for the valuation of office buildings as well as for the optimal exercise of the option to develop additional office space. Our empirical analysis concentrates on the latter. That is, we investigate the effects of explanatory variables suggested by these models on the trigger point at which additional office space should be developed. For example, if the trigger point increases in response to, say, an increase in the value of a particular explanatory variable, we should observe less development as compared to when the value of this variable is lower. In other words, we proxy movements in the trigger point by the number of buildings started, with a lower trigger point corresponding to a greater number of buildings starts.

We now turn our attention to the variables, both independent as well as dependent, relied upon in our subsequent empirical analysis. Their corresponding sample statistics are provided in Table 1.

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contrast, are less appealing and are generally deficient in floor plans, condition and facilities. Class C buildings are older buildings that offer basic services and rely on lower rents to attract tenants.

[7] Data are available only after the first quarter of 2000 for the remaining metropolitan areas included in the CoStar Office Report; we exclude them from our analysis because of the limited number of observations.

**Table 1** ■ Sample statistics.
This table presents various sample statistics for the independent and dependent variables used in our empirical analysis. We separately provide statistics for the sample of class A buildings versus the sample of class B buildings. Our data are quarterly in frequency over the 1998:III to 2002:II sample period and include the following U.S. metropolitan areas: Atlanta (ATL), Baltimore (BAL), Boston (BOS), Charlotte (CHA), Chicago (CHI), Cincinnati (CIN), Cleveland (CLE), Columbus (COL), Dallas-Fort Worth (DFW), Denver (DEN), Detroit (DET), Houston (HOU), Los Angeles (LAX), Northern New Jersey (NNJ), New York City (NYC), Orange County (LAO), Orlando (ORL), Philadelphia (PHI), Phoenix (PHX), Sacramento (SAC), San Diego (SDO), San Francisco (SFO), Seattle (SEA), St. Louis (STL), South Florida (SFL), Tampa (TAM), Washington D.C. (WAS) and Westchester County/Southern Connecticut (WCT).

| Variable | Mean | Median | Std Deviation | Minimum | Maximum |
|---|---|---|---|---|---|
| Class A buildings | | | | | |
| New building starts (per quarter) | 4.90 | 3 | 5.83 | 0: numerous cities | 37: SFO in 2000:1 |
| Log total number of buildings | 5.83 | 5.88 | 0.77 | 4.36: ORL in 1998:III | 7.14: LAX in 2002:2 |
| Lease growth rate | 0.05 | 0.05 | 0.03 | −0.01: NYC | 0.11: ORL |
| Vacancy rate (lagged) | 0.13 | 0.12 | 0.03 | 0.04: SFO in 2001:1 | 0.23: PHX in 2002:2 |
| Lease volatility | 0.08 | 0.07 | 0.04 | 0.03: CIN | 0.18: SFO |
| Class B buildings | | | | | |
| New building starts (per quarter) | 7.58 | 4 | 10.82 | 0: numerous cities | 94: PHX in 2000:1 |
| Log total number of buildings | 7.13 | 7.23 | 0.56 | 5.94: COL in 2000:I | 8.10: LAX in 2002:2 |
| Lease growth rate | 0.02 | 0.01 | 0.01 | −0.01: NYC | 0.04: DFW |
| Vacancy rate (lagged) | 0.11 | 0.12 | 0.03 | 0.05 SEA in 2000:3 | 0.20: DFW in 2002:2 |
| Lease volatility | 0.07 | 0.06 | 0.03 | 0.03: SFL | 0.13: SAC |
| 10-year real treasury yield (percent) | 2.72 | 2.63 | 0.86 | 1.52 | 4.20 |
| Term spread (percent) | 1.06 | 0.83 | 1.32 | −0.77 | 3.65 |
| Herfindahl ratio | 0.09 | 0.08 | 0.04 | 0.01: ORL | 0.18: TAM |

*Dependent Variable*

Our dependent variable is the number of buildings started during a quarter in a particular metropolitan area.[8] Unfortunately, this information is not directly tabulated in the CoStar Office Report. It can, however, be determined from data available on the number of buildings delivered and the number of buildings under construction. In particular, the number of starts during a quarter can be defined as the change in the number of buildings under construction plus the number of buildings delivered during the quarter.[9]

From Table 1 we see that the average number of class A building starts per quarter in a sampled metropolitan area is 4.90 with a median of 4 starts. The number of class B building starts is slightly higher with an average of 7.58 and a median of 4 starts. Since the number of building starts can only take on nonnegative values, the minimum number of starts observed is zero. The maximum number of class A buildings starts is 37 in San Francisco during the first quarter of 2000, and a maximum of 94 class B building starts is observed in Phoenix during the fourth quarter of 2000.

*Independent Variables*

The explanatory variables are chosen to correspond to variables suggested by the Grenadier and other real option pricing models with competitive interactions. We also include other explanatory variables in an attempt to address some of the previously discussed limitations of these models.

1. *Herfindahl ratio*: As of the end of our sample, the second quarter of 2002, the CoStar Office Report provides data on each metropolitan area's 10 largest developers ranked by the square footage of commercial real estate they developed in the preceding 12 months. These data are not disaggregated by building class as developers are not restricted to developing just one particular class of buildings. To measure the degree of competition prevailing in each region, we calculate a Herfindahl ratio defined as the sum of the squares of the market shares developed by each of the top five developers.[10] The larger this ratio, the more concentrated and so the less competitive a particular market.

   Rather than rely on each region's exact Herfindahl ratio in our empirical analysis, we reduce the noise inherent in this measure of competition by

[8] This is similar to Capozza and Li (2001) who analyze the number of single family home building permits issued.

[9] An obvious deficiency of this measure is that it treats all buildings as having the same size. The CoStar Office Report does not break down its data by building size.

[10] Our results are unaltered if instead we use data on all 10 developers.

forming a dummy variable which takes on the value of one if a particular market's Herfindahl ratio is greater than the sampled cities' mean concentration ratio and equals zero otherwise. We use this dummy variable to proxy for the level of competition, which in Grenadier's model is represented by the number of competitors. Since more competition is hypothesized to lead to more investment, we expect the coefficient on the Herfindahl dummy variable in a regression of new building starts to be negative.[11]

From Table 1 we see that the calculated Herfindahl ratios range from a low of 0.01 in Orlando, consistent with it being the most competitive market in our sample, to a high of 0.18 in Tampa, the least competitive market. Since the average of our sampled cities' Herfindahl ratios is 0.09, in our empirical analysis any market with a ratio greater than 0.09 is classified as being less competitive, *Herfindahl Dummy* $= 1$, while markets with a ratio of less than 0.09 are deemed to be more competitive, *Herfindahl Dummy* $= 0$.

2. *Growth of the lease rate*: For each metropolitan area, we have time series data on the corresponding quoted lease rates of class A and B buildings. The growth rate of the percentage change in the observed lease rates is calculated and used as a proxy for the demand shock's growth rate. Being a proxy for the drift of the demand shock process, we expect that the coefficient on the growth rate in a regression of new building starts to be positive. From Table 1, the average annualized growth rate in class A lease rates is 5% but varies from a low of $-1\%$ in New York City to a high of 11% in Orlando. Class B lease rates grew slower over our sample period, averaging only 2% with a low once again of $-1\%$ in New York City and a high of 4% in Dallas-Fort Worth.

3. *Volatility of the lease rate*: The volatility of the percentage change in these lease rates from one quarter to the next is calculated and used as a proxy for the volatility of the demand shock process. This variable is consistent with Grenadier's model save for the fact that in his model the lease rate is subject to a reflecting barrier corresponding to the case where development occurs. As a result, this may introduce a downward bias to our volatility estimates. However, in practice there does not exist a fixed boundary at which development occurs, so this bias should be minimal. As this is an option-based model, increases in volatility will

[11] Unfortunately we only have these data at a single point in time. Ideally we would like to have corresponding time series data. However, to the extent that the degree of competition changes slowly and the sample period is brief, our variable should provide a fairly accurate measure of the degree of competition prevailing in each market.

delay investment, and so we expect that the coefficient on volatility in a regression of new building starts should be negative.

For our sampled metropolitan areas, the mean annualized volatility of class A lease rates is 8% with a median of 7%, while the mean annualized volatility of class B lease rates is 7% with a median of 6%. The minimum annualized volatility of class A lease rates is 3% in Cincinnati, while San Francisco has the highest annualized volatility of class A lease rates of 18%. For class B lease rates, the minimum volatility is 3% in South Florida with a maximum of 13% in Sacramento.

Of course, a central message of the Grenadier and other real option pricing models with competitive interactions is that the degree of competition prevailing in a market should systematically influence the extent to which increases in volatility actually delay commercial real estate investment. In particular, the more competitive a market, that is, the smaller a metropolitan area's Herfindahl ratio, the shorter the investment delay that should be observed for a given increase in volatility. This implies a negative coefficient in a regression of new building starts on a term interacting volatility with our Herfindahl dummy variable.

4. *Treasury bond rate*: At the beginning of each quarter, interest rate conditions are measured by prevailing Treasury bond rates. We rely primarily on the real 10-year Treasury rate since commercial real estate development tends to be a longer-lived investment. Real rates are obtained by deflating corresponding nominal rates using the previous quarter's realized rate of inflation based on the Bureau of Labor Statistics' Consumer Price Index for all urban consumers index (CPI-U).

5. *Term spread*: The term spread, measured by the difference between the 10-year and 1-year nominal Treasury yields, is also included. As well as providing an alternative measure of prevailing interest rate conditions, the term spread is a business cycle proxy, and its inclusion allows us to investigate whether business cycle conditions influence office building starts.

6. *Total number of buildings*: For each metropolitan area, we have data on the total supply of class A and B buildings. Since these metropolitan areas differ in size, we need to control for these differences. Larger markets should experience more construction, and so the coefficient on this variable in a regression of new building starts should be positive. According to the CoStar Office Report, Los Angeles is the largest market for both class A and B buildings. Orlando is the smallest market for class A buildings, while Columbus is the smallest market for class B buildings.

7. *Lagged vacancy rate*: The vacancy rate variable is defined as the percentage of total vacant space for class A or B buildings divided by their total existing inventory. As noted earlier, contrary to Grenadier's model, there are good reasons to expect vacancies to actually occur and their level to be negatively related to new building starts. We lag the vacancy rate to capture the fact that it typically takes time for new building starts to respond to changes in vacancy rates.

   The mean vacancy rate for class A buildings is 13% with a median of 12%, being lowest (4%) in San Francisco during the first quarter of 2001 and highest (23%) in Phoenix during the second quarter of 2002. For class B buildings, the mean vacancy rate is 11% with a median of 12%. The minimum class B vacancy rate is 5% in Seattle during the third quarter of 2000, and the maximum is 20% in Dallas/Fort Worth during the third quarter of 2002.

## Empirical Method and Results

### *Poisson Regression*

The number of building starts is a count variable and only takes on nonnegative integer values. Being a count variable, there are numerous instances when we observe either zero or a very few starts per quarter; for example, over 50% of our sample of class A building starts involves no more than two starts per quarter. By contrast, there are only a few instances where we observe many starts during a quarter. While the number of building starts in a quarter can very well be zero, there is no *a priori* upper bound on the number of building starts that can be observed.

If $y$ denotes our count variable and $\underline{\mathbf{x}}$ is a vector of explanatory variables, we are interested in estimating the population regression, $E(y \mid \underline{\mathbf{x}})$. The most straightforward approach is a linear model, $E(y \mid \underline{\mathbf{x}}) = \beta\underline{\mathbf{x}}$ and estimating the parameter vector $\beta$ using ordinary least squares (OLS). Unfortunately, if $\hat{\beta}$ is the OLS estimator, there can be values of $\underline{\mathbf{x}}$ such that $\hat{\beta}\underline{\mathbf{x}} < 0$, so that the predicted number of new buildings started will be negative, clearly inappropriate for count data.

Alternatively, we will use a Poisson regression model to analyze our count data. That is, $y$ given the covariates $\underline{\mathbf{x}} \equiv (x_1, x_2, \ldots, x_k)$ is assumed to have a Poisson distribution whose density is

$$f(y \mid \underline{\mathbf{x}}) = \exp[-\mu(\underline{\mathbf{x}})][\mu(\underline{\mathbf{x}})]^y / y! \quad y = 0, 1, 2, \ldots \tag{12}$$

where $\mu(\underline{\mathbf{x}})$ denotes the conditional mean $\mu(\underline{\mathbf{x}}) \equiv E[y \mid \underline{\mathbf{x}}]$.

Assuming $\mu(\underline{\mathbf{x}}) = \exp(\underline{\mathbf{x}}\,\beta)$, which ensures positivity for any value of $\underline{\mathbf{x}}$ and any parameter value, as well as a random sample $\{(\underline{\mathbf{x}}_i, y_i)\colon i = 1, \ldots, n\}$, the log likelihood for observation $i$ is

$$\begin{aligned} \ell_i(\beta) &= y_i \log[\mu(\underline{\mathbf{x}}_i;\beta)] - \mu(\underline{\mathbf{x}}_i;\beta) \\ &= y_i \underline{\mathbf{x}}_i \beta - \exp(\underline{\mathbf{x}}_i \beta). \end{aligned} \tag{13}$$

The parameters of this model are easy to interpret. Since

$$\frac{\partial E(y \mid \underline{\mathbf{x}})}{\partial x_j} = \exp(\underline{\mathbf{x}}\beta)\beta_j$$

then

$$\begin{aligned} \beta_j &= \frac{\partial E(y \mid \underline{\mathbf{x}})}{\partial x_j} \times \frac{1}{E(y \mid \underline{\mathbf{x}})} \\ &= \frac{\partial \log[E(y \mid \underline{\mathbf{x}})]}{\partial x_j}. \end{aligned} \tag{14}$$

That is, $100\beta_j$ is the semi-elasticity of $E(y \mid \underline{\mathbf{x}})$ with respect to $x_j$. That is, for small changes in $x_j$, $\Delta x_j$, the percentage change in $E(y \mid \underline{\mathbf{x}})$ is approximately $100\beta_j \times \Delta x_j$.

Unfortunately, the Poisson model implies the equality of the conditional mean and variance of $y$:

$$\operatorname{var}(y \mid \underline{\mathbf{x}}) = E(y \mid \underline{\mathbf{x}})$$

which is often violated in practice. We will make the weaker assumption that

$$\operatorname{var}(y \mid \underline{\mathbf{x}}) = \xi^2 E(y \mid \underline{\mathbf{x}})$$

where $\xi^2 > 0$ is the variance-mean ratio to be estimated.

Finally, if $y_i$ given $\underline{x}_i$ is not Poisson distributed, then the estimator that maximizes the log likelihood function, Expression (13), is a quasi-maximum likelihood estimator. Under mild regularity conditions, the quasi-maximum likelihood estimator retains a number of desirable properties.[12]

[12] In particular, the quasi-maximum likelihood estimator is consistent and is efficient in the linear exponential family of distributions. For example, it is more efficient than the nonlinear least squares estimator. For further details, see Wooldridge (2002), especially Chapter 19, pages 645–683.

*Empirical Results*

Table 2 presents our estimation results. We consider a number of alternative regression specifications in the hope of better understanding the determinants of a developer's decision to start an office building.[13]

For each regression specification, we present the estimated coefficients as well as bootstrapped (in parentheses) $p$-values testing the null hypothesis that the corresponding coefficient equals zero. Bootstrapping[14] ensures that our inference is valid for finite sample sizes and accommodates other deviations from the maintained distributional assumptions that may be present in our data. For each regression specification we also tabulate the estimated dispersion, $\hat{\xi}$, and assess its fit using a *pseudo* $R^2$ statistic calculated as the squared correlation between the predicted and actual number of building starts.[15]

Regardless of the regression specification, a number of patterns emerge from the results of Table 2. For example, the expected rate of growth in demand proxied by the lease growth rate significantly influences the number of building starts. The higher this growth rate, the larger the number of class A and B building starts. As expected, we see that an increase in lagged vacancy rates tends to decrease the number of building starts. This effect is statistically significant in the case of class B building starts but not class A building starts. Also, we see that the larger the market, as measured by the log of the total number of buildings, either class A or class B, the greater the number of corresponding building starts.

Interestingly, we see across regression specifications and building classes that increases in the 10-year real Treasury yield are associated with a significant increase in the number of building starts. This is consistent with Capozza and Li (2002) who argue that real estate investment can increase in response to interest rate increases. For example, relying on the regression coefficients for class A

[13] While we rely on panel data, that is, data across different metropolitan areas and across time, fixed effects are captured in our Poisson regression by the fact that some of the explanatory variables—lease volatility, lease growth rate and Herfindahl dummy—are specific to a particular metropolitan area and do not change over time. By contrast, other explanatory variables—real T-bond rate and term spread—change over time and are not specific to a particular metropolitan area.

[14] To do so, we sample *with* replacement from the original data series and estimate a particular model using Poisson regression. We repeat this procedure 500 times to form an empirical distribution of each parameter of interest.

[15] Because this statistic does not mean what the $R^2$ statistic means in ordinary least squares (OLS) regression—the proportion of variance explained by the posited regressors–it should be interpreted with caution.

**Table 2** ■ Empirical results.
This table presents the results of estimating Poisson regressions of office building starts using various explanatory variables. The results are tabulated separately for class A versus class B buildings. The sample period begins in 1998:III and ends in 2002:II giving $n = 354$ observations. For each regression specification we provide corresponding estimated regression coefficients together with bootstrapped $p$-values (in parentheses) testing the null hypothesis that the coefficient equals zero as well as the regression's estimated dispersion, $\hat{\xi}$, and *pseudo* $R^2$ statistic, calculated as the squared correlation between the predicted and actual number of building starts.

| | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Class A buildings | | | | | |
| Intercept | −5.97 | −5.86 | −5.35 | −5.51 | −5.69 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Lease volatility | −2.11 | | −2.29 | −2.17 | −2.19 |
| | (0.03) | | (0.04) | (0.04) | (0.02) |
| Lease growth rate | 19.83 | 18.70 | 18.12 | 18.51 | 18.86 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Lagged vacancy rate | −2.13 | −0.60 | −2.22 | −2.16 | −2.00 |
| | (0.14) | (0.38) | (0.11) | (0.11) | (0.13) |
| 10-year real T-bond rate | 0.55 | 0.53 | 0.54 | 0.54 | 0.55 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Term spread | −0.46 | −0.47 | −0.46 | −0.46 | −0.46 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Total number of buildings | 0.96 | 0.91 | 0.89 | 0.91 | 0.93 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Herfindahl dummy | | −0.24 | −0.27 | | |
| | | (0.09) | (0.07) | | |
| Lease volatility × Herfindahl dummy | | | | −2.96 | |
| | | | | (0.15) | |
| Real T-bond rate × Herfindahl dummy | | | | | −0.07 |
| | | | | | (0.14) |
| $\hat{\xi}$ | 0.87 | 0.87 | 0.86 | 0.87 | 0.87 |
| pseudo $R^2$ | 0.51 | 0.50 | 0.51 | 0.51 | 0.51 |
| Class B buildings | | | | | |
| Intercept | −1.85 | −3.86 | −1.54 | −1.50 | −2.12 |
| | (0.05) | (<0.01) | (0.10) | (0.09) | (0.03) |
| Lease volatility | −8.84 | | −9.19 | −8.96 | −8.90 |
| | (<0.01) | | (<0.01) | (<0.01) | (<0.01) |
| Lease growth rate | 49.28 | 50.26 | 48.71 | 48.29 | 51.12 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Lagged vacancy rate | −11.62 | −2.70 | −7.99 | −8.04 | −9.14 |
| | (<0.01) | (0.21) | (0.02) | (0.02) | (0.01) |
| 10-year real T-bond rate | 0.35 | 0.34 | 0.34 | 0.34 | 0.37 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Term spread | −0.22 | −0.29 | −0.25 | −0.25 | −0.25 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Total number of buildings | 0.56 | 0.65 | 0.48 | 0.48 | 0.56 |
| | (<0.01) | (<0.01) | (<0.01) | (<0.01) | (<0.01) |
| Herfindahl dummy | | −0.42 | −0.44 | | |
| | | (0.01) | (<0.01) | | |
| Lease volatility × Herfindahl dummy | | | | −8.24 | |
| | | | | (0.03) | |
| Real T-bond rate × Herfindahl dummy | | | | | −0.15 |
| | | | | | (0.02) |
| $\hat{\xi}$ | 0.93 | 1.09 | 0.95 | 0.96 | 0.95 |
| pseudo $R^2$ | 0.28 | 0.29 | 0.31 | 0.31 | 0.31 |

building starts, all else being equal, the number of starts increase by approximately 0.55% for a one-basis-point increase in the 10-year real Treasury yield. Finally, the negative coefficient on the term spread across regression specifications and building classes implies that fewer building starts are observed when the term spread is wide. Recall that wide term spreads are typical of business cycle troughs. Conversely, narrow term spreads that characterize business cycle peaks are associated with more building starts.

Turning our attention to the alternative regression specifications themselves, Regression (1) relates office building starts to lease volatility irrespective of the prevailing level of competition. Consistent with the real options approach, we see that higher volatility, all else being equal, is associated with significantly fewer starts for both class A and B buildings. Intuitively, the higher the volatility, the more valuable is the option to delay, resulting in fewer starts. However, if volatility risk is priced in capital markets, this result is also consistent with higher volatility leading to a higher cost of capital which, in turn, decreases building values and results in fewer project acceptances. This risk-adjusted discounted cash flow perspective is in contrast to the real options approach where, all else being equal, a project's value is enhanced by higher volatility. Unfortunately, data on individual project values are unavailable to us and so cannot be relied upon to discriminate between these alternatives.

Regression (2) investigates the effect of competition on office building starts irrespective of volatility. Recall that competition is measured by a Herfindahl dummy variable, which takes on the value of one if a particular market's Herfindahl concentration ratio is greater than the sampled cities' mean concentration ratio and equals zero otherwise. This regression reveals that the more concentrated developers are in a particular market, all else being equal, the fewer the number of building starts. This relation is statistically significant at the 10% level for class A buildings and at the 5% level for class B buildings. Intuitively, the more concentrated a market, then, consistent with Grenadier, the less likely a developer's option to delay will be preempted. This result, however, is also consistent with standard microeconomic theory in which a monopolist sells a smaller quantity of a good than would a firm in a purely competitive market.

To ensure the robustness of these results, the simultaneous effect of volatility and competition on building starts is investigated in Regression (3). Notice that the magnitudes of these sensitivities as well as their statistical significance remain virtually unchanged when both lease volatility and the Herfindahl dummy variable are included.

Regression (4) attempts to discriminate between the real options approach and the risk-adjusted discounted cash flow model by relying on the fact that the interaction between volatility and competition should play a significant role in the former but not in the latter. For example, in Grenadier's real options model, the importance of competition as a determinant of building starts increases with increasing volatility. Competition only matters if volatility is high and developers fear the preemption of a valuable option to delay. At the other extreme, when volatility is absent and the option to delay is worthless, then the nature of competition has no effect on the observed number of building starts.

From Table 2 we see that the interaction between lease volatility and the Herfindahl dummy enters the regressions with a negative sign for both class A and B buildings, significantly so in the case of class B buildings. In other words, the delay in building starts brought about by an increase in volatility is attenuated by the presence of more competition among a market's developers. To gauge the economic significance of this interaction, the regression estimates for class B buildings, for example, give that the decrease in the number of building starts expected in response to a 1% increase in lease volatility is, all else being equal, approximately 8% greater if the development market is concentrated (*Herfindahl Dummy* $= 1$) then if it is not (*Herfindahl Dummy* $= 0$). This conclusion stands in contrast to the risk-adjusted discounted cash flow model in which the effect of volatility on building starts does not depend on the prevailing level of competition.

Pursuing this estimation strategy further, the real options approach suggests that a change in *any* underlying variable which increases the option value of a building project will make preemption by other developers more likely and so will attenuate the resultant delay in its development. If so, we should also observe that building starts are significantly affected by the interaction between the Herfindahl dummy and variables other than volatility that affect option value.

An increase in interest rates, for example, increases call option values, and so Regression (5) investigates whether the effect of interest rates on building starts is attenuated by the presence of more competition among developers. From Table 2 we see that this indeed is the case for both class A and B buildings with the interaction between the 10-year real T-bond rate and the Herfindahl dummy having a statistically significant effect on the observed number of class B building starts. In particular, while an increase in interest rates is still associated with an increase in the number of class B building starts, this response is almost halved in markets where developers are more concentrated.

## Conclusion

Real options models provide economists with a better understanding of investment decision making under uncertainty. Researchers have recently begun to investigate the role of competition in these models. This article empirically investigates the importance of competition in actual commercial real estate investment decisions. We find that the number of office building starts across 28 U.S. metropolitan areas over the 1998 to 2002 sample period is indeed influenced by the competitive nature of the local commercial real estate market. Consistent with Grenadier (2002, 2005) and other real option pricing models, competition has a significant effect on commercial real estate development by interacting with volatility to attenuate the value of the developer's option to wait.

Empirical tests such as those provided in this article are also important in identifying the deficiencies of current real options models and suggesting avenues for future research needed to improve their applicability. For example, we find that volatility plays an important role in determining building starts beyond its interaction with competition. Whether this result reflects our use of the Herfindahl ratio to characterize competition or continues to hold when other measures of competition, such as that relied on by Bulan, Mayer and Somerville (2002), are used should be explored in future research.

Incorporating game theoretic concepts into real options models clearly improves our understanding of commercial real estate investment decisions. It should also improve our understanding of investment decision making in other markets as well.

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### Queries

**Q1** Author: Please update reference Bulan et al. (2002).
**Q2** Author: Please update reference Novy-Marx (2002).