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James D. Shilling
University of Wisconsin - Madison

C. F. Sirmans
University of Connecticut

John B. Corgel
Cornell University School of Hotel Administration, jc81@cornell.edu

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Natural Office Vacancy Rates: Some Additional Estimates

Abstract

[Excerpt] In our recent paper (Shilling et al. [2]), we examined the rent adjustment process for 17 U.S. office markets and provided the first estimates of the natural vacancy rates. In his comment, Voith [3] argues that our specification of the rent-vacancy relationship is biased since we introduced an interaction term to capture the effect of risks associated with higher vacancy levels. He suggests that an alternative specification would be to enter a vacancy squared term to capture this risk effect.

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rent-vacancy relationship, risk, vacancy levels

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Natural Office Vacancy Rates: Some Additional Estimates

James D. Shilling
University of Wisconsin-Madison

C.F. Sirmans
University of Connecticut

John B. Corgel
Cornell University

In our recent paper (Shilling *et al.* [2]), we examined the rent adjustment process for 17 U.S. office markets and provided the first estimates of the natural vacancy rates. In his comment, Voith [3] argues that our specification of the rent-vacancy relationship is biased since we introduced an interaction term to capture the effect of risks associated with higher vacancy levels. He suggests that an alternative specification would be to enter a vacancy squared term to capture this risk effect.

To examine the validity of his comments, we reestimated our equations using three specifications of the rent-vacancy relationship:

$$\dot{R} = b_0 + b_1 E - b_2 V + e \tag{1}$$

$$\dot{R} = \alpha_0 + \alpha_1 E - \alpha_2 V + \alpha_3 (\dot{R})V + u \tag{2}$$

$$\dot{R} = \beta_0 + \beta_1 E - \beta_2 V + \beta_3 V^2 + w. \tag{3}$$

Resulting values of V^n are reported in columns 1-3 of Table 1. The mean values of the natural vacancy rate range from a low of 3.91% to a high of 8.59%.[1] Results from (1) yield negative and hence implausible natural vacancy rates for New York City and San Francisco. Estimates from (2) are virtually unchanged when a lagged interaction term is used. The mean level of V^n is 8.03% when the lagged interaction term is used and the Pearson coefficient of correlation between the estimates from (2)

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[1] All equations were estimated using ordinary least-squares. The normal vacancy rate for each city is determined from the regression results by assuming $V^n = b_0/b_2$

TABLE 1

Estimates of Natural Vacancy Rates in Commercial Office Buildings for Selected Cities under Alternative Model Specifications[a]

City	1 Without interaction term (%)	2 With interaction term (%)	3 With squared term (%)	4 Voith and Crone's estimates (%)	5 Wheaton's estimates (%)
Atlanta	10.0	6.3	6.6	17.6	13.2
Baltimore	5.9	13.9	1.9	2.5	—
Chicago	8.0	4.1	5.5	7.2	8.2
Cleveland	5.7	2.8	1.4	9.9	—
Denver	20.0	12.3	4.6	14.6	8.3
Des Moines	5.1	9.9	4.1	—	—
Detroit	6.7	11.8	3.3	—	—
Indianapolis	7.6	6.5	5.4	—	—
Minneapolis	1.4	4.5	3.5	—	—
New York City	—	1.0	0.2	5.8	6.1
Philadelphia	3.2	9.5	6.5	8.3	8.7
Pittsburgh	8.2	10.0	3.4	5.7	—
Portland	23.9	16.0	4.7	—	—
San Francisco	—	2.9	1.9	—	6.9
Seattle	4.9	8.4	4.0	—	—
Spokane	9.7	10.5	5.7	—	—
Mean	8.59	8.15	3.91	8.95	8.57

[a]Models 1, 2, and 3 use data for 1960–1975; Voith and Crone use data for 1979–1987; Wheaton uses data for 1980–1988.

and those when the lagged interaction term is used is 0.56. We interpret these results to suggest that the criticism raised by Voith [3] may be unwarranted. Furthermore, the pattern of residuals from (3) when plotted against the vacancy variable does not indicate the need for a quadratic term.

A Duncan's multiple range test suggests that the difference between the mean natural vacancy rates obtained in (1) and (2) can be attributed to random sampling fluctuations. The test further suggests that the mean natural vacancy rates from (1) and (2) are different from the mean natural vacancy rate for (3). The Pearson coefficient of correlation between the natural vacancy rates in (1) and (2) is 0.57. Thus it would appear that our original estimates are consistent with Rosen and Smith's [1] basic theoretical model of the rent-vacancy relationship. On the other hand, the Pearson coefficient of correlation between the values of V^n in (1) and (3) is 0.33, and between (2) and (3) is 0.20. These

results indicate that the alternative specification suggested by Voith may actually lead to statistically different results.

In Table 1, we also show estimates of the natural office vacancy rate from Voith and Crone [4] for the 1979 to 1987 period, and Wheaton [5] for the 1980 to 1989 period. The mean V^n for these two equations are 8.95% and 8.57%, respectively. A Duncan's multiple range test for those cities with available data suggests that the difference between the mean natural vacancy rates in columns 1, 2, 4, and 5 are different from the mean natural rate in column 3.

The Pearson coefficients of correlation between our estimates and Voith and Crone is zero, whereas the correlation between our estimates and Wheaton's [5] is 0.39. These comparisons may be misleading since Voith and Crone's [4] model does not control for the relationship between changes in rents and deviations from natural vacancy rate. Our recent paper [2, p. 99] suggests that the price-inventory adjustment process for commercial office space depends heavily on deviations in the actual vacancy rate from the desired vacancy rate: "[r]eactions of output and prices to demand changes were strongest when the gap between desired and actual inventory holdings was the largest." We also found that "differences in the marginal costs of carrying inventory... help to explain variations in the normal vacancy rate" (p. 91). Wheaton's [5] estimates are obtained from a model similar to Eq. (1).

A second factor confounding the comparisons in Table 1 is the different estimation periods. We showed in our earlier paper that variations in the natural vacancy rate can be explained by differences in variables measuring expected growth in demand and supply of office space, and the marginal costs of holding inventories. This makes the estimation of the natural vacancy rate susceptible to structural change, a subject about which we know very little. Anecdotal evidence, for example, suggests that the composition of the national office market has changed in the last 15 years with large downtown centers shrinking in importance and demand. This structural change follows a trend set by large office tenants who seem to be moving from costly urban space to less expensive suburban space. It also appears that the office market has been overbuilt for some time. In some parts of the country, notably the Northeast and the Oil Belt in the Southwest, the new groundswell is a

slow-growth movement. In other parts of the country, regional infrastructure and higher tenant turnover seem to be contributing to a slower rate of increase in new office space and higher vacancy rates. Another untested hypothesis is whether recent tax law changes have lowered the natural vacancy rate for office space.

Modeling these structural changes should prove extremely crucial to the estimation of the rent adjustment equation. The robustness of our results to alternative specifications of the rent-vacancy relationship would further suggest that accounting for structural change may be more important to the estimation of the natural office vacancy rate than the method used to capture the effect of risks associated with higher vacancy levels.

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