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Capitalization of Property Taxes: Vacant Properties

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1. Introduction

A principal proposition of public finance theory is that a discriminatory tax on an asset will be capitalized into its market value. A number of empirical studies have attempted to verify this proposition.[1] However, the methods used in these tests have been criticized by Wales and Wiens [1974].[2] They show that the use of the effective tax as a determinant of the market value of property introduces a spurrious correlation between market value and the effective tax rate which biases the empirical tests in favour of the capitalization hypothesis.

The purpose of this paper is to test for capitalization of property taxes using a nonlinear model derived from the formal theory of capitalization and to compare the results with those yielded by the test developed by Wales and Wiens. To our knowledge, the nonlinear model has not been used in any previous study.

The nonlinear model indicates that property taxes are capitalized to a small extent, while the Wales-Wiens test indicates that the null hypothesis of no capitalization cannot be rejected. Ostensibly the two tests yield contradictory results. However, given the biases in both models we conclude that property taxes are not capitalized, or at best they are capitalized to a small degree.

2. The Wales-Wiens Test

In this paper we attempt to test the extent to which differences in property taxes are capitalized into the market value of property in a single tax jurisdiction. Taxes may differ on virtually identical properties in a given jurisdiction because a random element exists in the determination of assessment values.[3] The present value of any given property is

$$V = \sum_{t=0}^{N} (Y_t - T_t \cdot \gamma)/(1+r)^t \tag{1}$$

where N is the number of years the investor expects to hold the property, Y_t is the expected return in year t, r is the relevant rate of interest, T_t is the expected tax payment in year t, and γ ($0 \leq \gamma \leq 1$) reflects the degree of capitalization. We hypothesize that the expected return on a property is

$$Y_t = X \cdot B_t \tag{2}$$

where X is a vector of property characteristics and B_t is a vector of imputed values attached to the characteristics.

If we assume that property taxes and the market conditions underlying B are constant then the following linear approximation of equation (1) can be used to test for capitalization:

$$V = XB_1 + T\gamma_1 + \mu_1 \tag{3}$$

where μ_1 is a random disturbance with the usual properties. When Wales and Wiens [1974] and King [1972] estimated equation (3) they found the tax coefficient γ_1 to be positive.[4] The difficulty with using equation (3) is that due to lack of data some property characteristics might be excluded from X. If these omitted variables are correlated in the same direction as both V and T then the estimate of γ_1 will be positively biased.[5] If the omitted variables are important determinants of market and assessed values, their influence may be strong enough to bias γ_1 upwards sufficiently that it will be positive, even in the event of capitalization.

Wales and Wiens suggest that the possible presence of bias due to omitted variables can be eliminated by replacing T in equation (3) by the effective tax rate

$$\bar{m} = T/V \tag{4}$$

Thus instead of estimating equation (3) they suggest that the following equation might be more appropriate:

$$V = XB_2 + (T/V)\gamma_2 + \mu_2 \tag{5}$$

The rationale is that the inclusion of V on the right-hand side of the regression equation might cancel the major effects of the omitted variables on T and V. However, while the omitted variable bias might be eliminated, a new bias will be introduced if some of the stochastic variation is due to factors other than omitted variables. For example, differences in tastes may result in varying selling values with properties with an identical set of X characteristics. But differences in tastes do not affect tax payments. Thus the estimate of γ_2 might be negative, and possibly significant, even in the event of no capitalization.

If the capitalization effect can be separated from the negative bias induced by the inclusion of V on the right hand side of (5) a valid test for capitalization is possible. Wales and Wiens have devised a method by which the two effects may be separated. Essentially, they attempt to estimate the bias.

In order to estimate the bias they first attempt to purge V of any capitalization effect and then use both V and the purged value V' in the regression equation. To arrive at V' market value is regressed on property characteristics

$$V = XB_3 + \mu_3 \quad (6)$$

to obtain estimates $\hat{B}_3$ and $\hat{\mu}_3$ of B_3 and μ_3 respectively. Then the new series V' is generated as follows:

$$V' = XB_3 + \mu_4 \quad (7)$$

where μ_4 is generated from a normal distribution with mean zero and standard deviation $\hat{\mu}_3$. V in equation (5) is then replaced by V'.

$$V' = XB_4 + (T/V')\gamma_4 + \mu_5 \quad (8)$$

to obtain estimates for γ_4 and B_4.

Since V' is generated independently of any tax effects, in the event of no capitalization equations (5) and (8) will be based on the same underlying population and thus yield estimates which are not significantly different from one another. In the event of no capitalization γ_4 can be interpreted as a measure of the bias induced by the inclusion of V on the right hand side of equation (5). Unfortunately, in the event of capitalization the estimate of γ_4 will be biased in an indeterminate direction.

In the event of capitalization the following biases occur. First, the estimate of the variance of μ_3 (used to generate V') will be biased upwards. This tends to bias the estimate of γ_4 in a negative direction. Second, the OLS estimate of each component of B_3 will be biased in a direction determined by two factors[6] - the correct sign of γ_2 and the sign of the correlation between the effective tax rate and the associated property characteristic. If γ_2 and the correlation have the same sign, then the component of $\hat{B}_3$ will be biased upwards. Otherwise, it will be biased downwards. In the event of capitalization, γ_2 is negative; however, the correlation matrix indicates that the effective tax rate is positively correlated with some variables, negatively correlated with others. Thus some of the estimated coefficients will be biased upwards, and some downwards. As a result the net effect of the biases is indeterminate.

The above discussion implies that we might not reject the null hypothesis even though capitalization actually occurs. If the null hypothesis is rejected, we are still left without an unbiased estimate of the degree of capitalization. However, if the error induced by the differing biases in equations (5) and (8) is small relative to that induced by estimating equations (3) or (5), then $(\gamma_2-\gamma_4)$ may be a more satisfactory measure of the degree of capitalization than the estimate of either γ_1 or γ_2.

3. The Nonlinear Test

Let us now consider an alternative to the Wales-Wiens test. In the last section we replaced T in equation (3) by $\bar{m} = T/V$. A more correct way to proceed would be to replace T in equation (1) by $\bar{m}V$. Then solving for V and using equation (2) we get

$$V = XB_5/((r/(r+1)) + \bar{m}\gamma_5) \qquad (9)$$

Equation (9) cannot be estimated by OLS because it is intrinsically nonlinear - that is, it is nonlinear in the parameter γ_5. Thus nonlinear least squares is used to obtain estimates for B_5 and γ_5.[7]

Because capital markets are not perfect, there is no single discount rate r which will give us a unique value for that parameter. Thus we postulate several different rates (each reflecting a different opportunity cost of investing in unimproved properties) and observe any effects that the varying rates might have on the parameter estimates. Note that if we assume investors start discounting at t=1 instead of t=0, the specification would be

$$V = XB_6/(r + \bar{m}\gamma_6) \qquad (10)$$

We found that the initial period of discounting does not have a significant effect on the coefficient estimates.

Equations (9) and (10) have the same problem as equation (5). The estimates of γ_5 and γ_6 will be biased downwards because of the inclusion of V on the right hand side of the regression equation.

4. Data

The sample data consists of information on 543 unimproved properties which were sold in the municipality of Surrey in 1973.[8,9] In addition to sale prices and tax payments, information is available on the following characteristics of each property: lot frontage (LOTFRNT), frontage rate (FRNTRATE), cash or noncash sale (CASH), and whether a building could have been constructed on the property immediately after the sale was made (UNBUILD). We also know in which quarter of the year each property was sold. A few comments on the property characteristics may be necessary for the reader who is unfamiliar with assessment procedures.

Lot frontage is the effective measure of frontage on the street after adjustments have been made for the lot's depth, shape and terrain. We would expect a positive relation between lot frontage and market value.

Frontage rate is the assessor's estimate of the average per foot value of the property on the street. It is based on past sales (the 1973 frontage rate is based on 1971 sales) and can be interpreted as a proxy measure for the quality of the neighbourhood. Clearly we would expect a positive relation between frontage rate and market value. CASH is a code which indicates whether cash was paid for the property or if there was an alternative financial arrangement made at the time of sale (for example, the seller might have given the buyer a mortgage); CASH enters the regression equation as a zero-one dummy - it equals one if cash was paid for the property, zero otherwise. We would expect a negative relation between CASH and market value, as a discount is generally available on cash sales. It is interesting to note that 78 per cent of the vacant lot sales in 1973 were cash sales. The percentage of cash sales is high because it is difficult to obtain a mortgage on unimproved properties, possibly because the purchase of such properties is considered more of a direct investment than the purchase of improved properties. UNBUILD is a code, also entering the regression as zero-one dummy - it equals one if a building cannot be put on the property immediately, zero otherwise. We would expect a property that is "buildable" to command a higher price on the market, and therefore hypothesize a negative relation between UNBUILD and market value. Approximately 36 per cent of the lots sold in 1973 were "unbuildable".

The price increase in the real estate market was very rapid in 1973. Thus sales values will be inflated according to the time of year the property was sold. But taxes are based on a predetermined assessment value, and therefore are not subject to the rise in prices. The implication is that the effective tax rate will decrease sequentially on properties as they are sold later in the year.[10] In terms of the original model (equation (1)) this means that Y_t and T_t are measured in constant dollars while V is measured in current dollars. We should use the deflated value of properties in the regression equations. Lacking appropriate deflators we initially attempt to circumvent the inflation problem by using quarterly sample data in the nonlinear test.[11]

However, resorting to quarterly estimation is not entirely satisfactory since the tax payments and returns are annual. Furthermore, we do not end up with a unique measure of the extent of capitalization. Therefore we will use annual data and test whether the extent of inflation is sufficient to cause structural change in our model between quarters.

5. Nonlinear Regression Results

The following equation was estimated for each of the four quarters

$$\ln V = B \cdot X - \ln(r/(r+1) + \bar{m} \cdot \gamma) + \mu \qquad (11)$$

By expressing equation (11) in logarithmic form we reduce the possibility of a heteroscedastic disturbance.[12] If μ, remains significantly heteroscedastic, the coefficient estimates will be unbiased, but there will be a loss of efficiency in the variances of the estimates; the implication is that we are less likely to reject the null hypothesis of no significance when we test for the statistical significance of the independent variables.

We experimented with five different discount rates in each quarter - 2%, 4%, 6%, 8%, and 10%. The first observation made was that as the discount rate is varied (in any given quarterly regression), the estimated coefficients of the property characteristics do not change significantly -

in most cases they do not change at all. However, we observe that the estimated coefficient of the effective tax rate (EFFTAX) does vary with the discount rate. As successively higher rates are used, the absolute value of the estimated coefficient increases. This is an important observation since the estimated coefficient of EFFTAX reflects the degree of capitalization. The implication of the observation is that the extent of capitalization is positively related to the opportunity cost of investigating in unimproved properties.[13] However, the degree of capitalization is not highly sensitive to the discount rate. We shall see shortly that the magnitude of variation between the estimated coefficients obtained when using the two extreme discount rates is small. In view of the above observations, in subsequent discussions we focus on the regression results obtained when using the two extreme discount rates.

First equation (11) was estimated including in X every property characteristic for which data is available. Almost all of the estimated coefficients except UNBUILD are consistently of sign and magnitude in line with *a priori* reasoning. The coefficients of UNBUILD are disconcerting. In the first and the third quarters the sign of the coefficient is positive, but in the second and fourth quarters, it is negative. Furthermore, the coefficient is statistically significant from zero in one quarter (third quarter), but not in the others. Because these results are nonsensical, we omitted UNBUILD from X and re-estimated equation (11), in the hope that both the regression coefficients of the other characteristics and the fit of the equation would not change significantly. In fact this is what we observe.[14] Thus in subsequent estimations we continue to omit UNBUILD from X.

The regression results obtained when UNBUILD was excluded from equation (11) are reported in Table 1. Examine first the non-dummy property characteristics, LNFRNT and LNFRATE. Both are highly significant in each quarter and both are large relative to the other coefficients. These observations indicate that the two variables are important determinants of market value. Since these variables and the independent variable are in logarithmic form, the estimated regression coefficients are interpreted as elasticities. Consider, for example, the estimated coefficients from the first quarter regression. The coefficient of LNFRNT is 0.81-- thus a 1 per cent increase in property size will raise property value by 81 per cent. Similarly, a 1 per cent increase in the frontage rate raises property value by 88 per cent (coefficient of LNFRATE is 0.88).

The estimated coefficients of the dummy variables are interpreted differently. Because the dependent variable is in logarithmic form while the dummies are not, the estimated coefficients are interpreted as percentages rather than as elasticities. After omitting UNBUILD, only one dummy, CASH, is present in the quarterly regression,[15] according to our estimates, if property bought say in the first quarter is paid for in cash, the investor will pay 3 per cent less for it than if a mortgage were obtained.

Upon scanning across quarters the coefficients and the t-statistics of the cash variable, we observe that CASH is significant only in the third quarter. Furthermore, the estimated coefficient is considerably larger in the third than in the other quarters. According to the third quarter estimate, if a property is sold for cash, the purchaser will pay 22 per cent less for it than if a mortgage were obtained - a curious result, one which we find difficult to explain.[16] The estimated coefficients for CASH in the other quarters, however, are consistent with a priori reasoning. Nevertheless, it is disconcerting that when the estimates are "reasonable", CASH is statistically insignificant, but when the estimate is not "reasonable", CASH is significant.

We next examine the estimated coefficient of EFFTAX which is of particular interest because it measures the extent of capitalization. As we have seen, the magnitude of the estimated coefficient of EFFTAX varies directly with the discount rate. If the discount rate is 2 per cent, the estimated coefficient of EFFTAX is 0.01 in the first and last two quarters and 0.02 in the second quarter. If it is 10 per cent, the estimated coefficient of EFFTAX is 0.04 in the first and last quarters, 0.05 in the third quarter, and 0.09 in the second quarter. Thus if the discount rate is 2 per cent, approximately 2 per cent of the property tax will be capitalized into the market value of the lot. Note that in all cases EFFTAX is statistically significant from zero.

Although the estimated coefficient of EFFTAX is sensitive to the discount rate, and in some cases to the quarter in which the property is sold, we can make two general conclusions - i) some capitalization does occur, for EFFTAX is statistically significant irrespective of the discount rate and the quarter, and ii) extent of capitalization is small - on average it does not exceed 5 per cent.

ii) Pooled Regressions

Upon examining the quarterly regression results, we observe that the coefficients of each independent variable appear to be stable[17] across quarters with two exceptions - the coefficient of EFFTAX in the second quarter, and that of CASH in the third quarter.[18] We hypothesize that the instability of these coefficients might be attributed to institutional pecularities that could have arisen in the real estate and/or financial market in the second and third quarters. However, in view of the general stability of the quarterly regression results, the data was pooled, and equation (11) re-estimated.[19] We then formally tested the homogeneity or stability of the relationship over tne four quarters; that is, we tested for the equality of the quarterly coefficients of each variable. We performed the test, because if we found the relation to be stable over the four quarters, we would have a more satisfactory annual model to work with.

The stability of our model over four quarters is tested by comparing the residual sum of squares from the annual regression with the addition of the residual sum of squares from the quarterly regressions. In the pooled or annual regression the quarterly coefficients are implicitly restricted to be equal, while in the four quarterly regressions, they are allowed to vary from quarter to quarter. Thus we would expect the residual sum of squares from the pooled regression (restricted residual sum of squares) to exceed the addition of the residual sum of squares from each of the quarterly regressions (unrestricted residual sum of squares). The equality of the coefficients is thus tested by comparing the difference between the restricted and unrestricted sum of squares.

The results obtained from the pooled regression are reported in Table II. Note the three quarterly dummies that have been added to the regression equation - YQ2, YQ3, YQ4 (these are the second, and fourth quarter dummies respectfully). The estimated coefficients of these dummies are interpreted as the quarterly rate of inflation in the price of unimproved properties, for they reflect the variation in sale price accounted for by the quarter in which the sale was made. These coefficients reflect "pure" price changes since they measure the change in price resulting from a change in quarter while holding property characteristics constant.[20] According to the estimates, prices were inflated by 2% in the second quarter, 10% in the third quarter, and 18% by the fourth quarter (note that owing to the method by which the quarterly dummies have been set up, the base quarter from which inflation is measured is the first quarter of 1973.) Only the second quarter dummy (YQ2) is statistically insignificant; since exclusion of YQ2 from the regression equation does not alter the other estimated coefficients nor reduces the fit, it is left in the equation because it is part of a quarterly price index for 1973.

The other estimated coefficients are of magnitude and sign consistent with a priori reasoning. Note that they are all statistically significant. In view of the quarterly regression results this observation is not surprising except for the case of CASH. In three of the quarterly regressions CASH is insignificant, but in the pooled regression it becomes significant.

Upon testing the equality of the quarterly slope coefficients, we found that the hypothesis of equality had to be rejected.[21]

Thus we set up an alternative pooled regression. Instead of constraining the coefficients of each variable to be equal in all quarters, we allowed the coefficient of CASH to vary in the third quarter, and that of EFFTAX to vary in the second quarter, for we have observed that these coefficients appear to be unstable. Then using the method described above we tested whether the quarterly regressions satisfy the same relation as the alternative constrained regression. That is, we tested whether the estimated coefficients of LNFRNT and LNFRATE are equal in all quarters, those of EFFTAX equal in the first, third and fourth quarters. The regression results are reproduced in Table III. We found that we could not reject the hypothesis that the constrained and unconstrained[22] regressions satisfy the same relation.

To confirm the test, we then tested the hypothesis that the constrained regressions satisfy the same relation - the hypothesis is rejected thereby confirming the previous test.[23]

The regression results in Table III are taken as the final estimates. Note that CASH is insignificant in the first, second and fourth quarters. Further note that the estimated coefficients of the quarterly dummies can no longer be interpreted as the rate of inflation since the quarterly dummies were employed to allow variance in the coefficients of CASH and EFFTAX. This is unfortunate for if we had estimates of the rate of inflation we could deflate market values and therefore eliminate the inflation problem by using the deflated values in the annual regression. The estimates from Table II cannot be used as deflators since the relation is not stable.

Of particular interest are the coefficients of the tax variable. We found that the coefficients of EFFTAX are equal in the first, third, and fourth quarters. In these quarters the coefficient is .05 for a 10% discount rate and .01 for a 2% discount rate. In the second quarter, the coefficient is .08 for a 10% discount rate, and .02 when the discount rate is 2%. Thus the extent of capitalization is greater in the second quarter, possibly due to an institutional factor; the tax coefficient is larger by .03 when the discount rate is 10%, and by .01 when it is 2%. We conclude that the extent of capitalization varies between five and eight per cent for a high discount rate, and between one and two per cent for a low discount rate. However, regardless of the variation stemming from the discount rate and from a possible institutional peculiarity in the second quarter, the extent of capitalization is small.

6. Wales-Wiens Regression Results

When the logarithmic versions of equations (4) and (5) were estimated, the significance of the tax variables and the sign of the regression coefficients (Table IV) were consistent with our expectations which were based on the possible bias problems discussed above: when market value was regressed on tax payments (equation A), the estimated coefficient of the tax variable came out positive (0.0370), but insignificant (t = 1.8419); when the effective tax rate was used in place of tax payments (equation B), the corresponding estimated coefficient came out negative (-0.3593) and highly significant (t = 14.6965). According to the last estimate, a one percentage change in the effective tax rate would depress market value by 35 per cent![24] R^2 for both equations is good for a cross-sectional study - 0.60 and 0.71 for equations A and B respectfully. But approximately one-third of the variation in prices is not explained in the regressions, thus leaving open the possibility that important variables have been excluded and/or that there are substantial taste differences among consumers. The implication is that it is highly unlikely that the regression coefficient of the tax variable from either equation A or B is an unbiased measure of the degree of capitalization, and that either t-statistic can appropriately be used to test the significance of the tax variable. Thus, in order to test the capitalization hypothesis, equations (4) to (8) were estimated in conformity with the Wales-Wiens methodology.

After estimating equation (8) (Equation C-Table IV) an F-Test[25] was performed on the regression coefficients of the tax variables in equations (5) and (8) in order to determine if they are significantly different. If the coefficients do not differ significantly, the implication is that the two equations are based on the same underlying population - thus the null hypothesis of no capitalization will not be rejected.

The F-Test is set up in the following manner. First an equation with $\ln V$ and $\ln V'$ as the dependent variable, and with two tax variables on the right-hand side - T_i/V or T_i/V', depending on whether the observations are from generated or actual data, and $(T_i/V).d$, where d = 1 if actual observation, and zero if generated observation - is estimated (equation D). Then we test the significance of $T_i/V.d$. Under the null hypothesis of no capitalization, the coefficient of $(T_i/V).d$ will not differ from zero since it reflects the effect of taxes on market value over that induced by the spurious correlation. In fact this is what we observe. Therefore we do not reject the null hypothesis of no capitalization.

7. Conclusions:

We have used two alternative tests for capitalization of property taxes into the market value of vacant properties. Although the nonlinear test rejects the null hypothesis of no capitalization, the degree of capitalization is limited. However the bias of this test favours rejection of the null hypothesis. On the other hand, when we used the Wales-Wiens test we could not reject the null hypothesis of no capitalization. However, the net effect of the biases in this test is indeterminate.

We can combine these two results by examining two cases. First, if the net effect of the biases in the Wales-Wiens test favours rejecting the null hypothesis, then we can unambiguously say that we could not reject the null hypothesis of no capitalization. Second, if the net effect favours not rejecting the null hypothesis, then the Wales-Wiens test may be wrong. Thus capitalization may occur. But given the nonlinear test, we conclude that the degree of capitalization is limited.

FOOTNOTES

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1. See Netzer [1966] and Aaron [1975] for summaries of results and Daicoff [1967], King [1972] [1977], Oates [1969], Rosen and Fullerton [1977], Smith [1966] and Wales and Wiens [1974].

2. Wales and Wiens tested for capitalization on residential properties in the municipality of Surrey. In this study we test for capitalization on vacant properties in the same municipality.

3. This fact is well known. See for example White and Hamilton [1972].

4. In footnote 9 Wales and Wiens mention that they estimated equation (3) in double log form since they expect changes in the effective tax rate to change V by different absolute amounts depending on the level of V. King [1977] has re-estimated Oates' [1969] equations taking this point into consideration.

5. This point was made by Wales and Wiens. The estimate of γ_1 will be positively biased if omitted variables are either negatively or positively correlated with both the dependent variable and the tax term. Since property taxes are a function of assessed values and assessed values are based on market values in some base period, both the tax term and the dependent variable will generally be correlated in the same direction with any omitted variable. King [1977], footnote 6, doesn't agree.

6. In the event of multicollinearity, matters are more complicated. The force of the argument remains the same.

7. See Draper and Smith [1966] chapter 10.

8. The data was made available by the District of Surrey.

9. Surrey is a municipality in the Greater Vancouver area which has a population of approximately 100,000.

10. We calculated the mean effective tax rate of each quarter to test this hypothesis. In line with our expectations, the mean rate fell successively from the first quarter to the last quarter - 1.5, 1.2, .97 and .84.

11. In the Wales-Wiens test the bias introduced by inflation is incorporated into the general bias problem. Thus there is no need to resort to quarterly regressions.

12. See footnote 4.

13. This is what we would expect. For instance, suppose the perpetual net income stream of an asset is $12. If r=.02, discounting from t=1, the market value of the asset is $600; if r=.04, market value is $300, and so on. As successively higher discount rates are used, the market value of the asset falls, that is, the degree of capitalization becomes greater (the estimated coefficient of EFFTAX increases).

An intuitive explanation is that the discount rate reflects the after-tax rate of return on other assets. We would expect a greater degree of capitalization to occur for a given net return on unimproved property the greater the return on other assets, for the market value of the property in line with that on other assets.

14. The results are not surprising since UNBUILD is statistically significant only in the third quarter. The fit of the equations remain unchanged (to two decimal places) when UNBUILD is dropped from the first, second and fourth quarter regressions, but it falls from 0.75 to 0.73 when dropped from the third quarter regression.

15. In subsequent pooled regressions, quarterly dummies are also present.

16. Our curiousity is further aroused by the fact that UNBUILD is also significant only in the third quarter. We suspect thst an unusual event was occurring in the third quarter which might explain these results.

17. The stability of the quarterly regression coefficients of a given variable is tested by examining the corresponding confidence intervals (95% level). Considerable overlap of the confidence intervals would indicate that the coefficient in question is stable.

18. This finding is expected since the coefficients of EFFTAX and CASH are relatively large in these quarters. It is not surprising that some of the coefficients of EFFTAX and CASH are unstable, for these variables are the ones most likely to be sensitive to the prevailing financial climate.

19. Three seasonal dummies were included in X_1 - these of course were not present in the quarterly regressions.

20. Thus a hedonic price index can potentially be constructed from these coefficients.

21. See appendix, Section II.

22. The term "unconstrained" is used for ease of exposition - we are actually referring to four quarterly regressions.

23. See appendix, Section III.

24. The coefficient is interpreted as a percentage since (T/V) is not expressed in logarithmic form, while V is.

25. See Wales and Wiens, footnote 19.

TABLE I

QUARTERLY REGRESSIONS - r=2% and r=10%

ESTIMATED COEFFICIENTS

	First Quarter	Second Quarter	Third Quarter	Fourth Quarter
		r=2%		
CONSTANT	-6.5911 (-23.29)	-6.1673 (-23.10)	-4.1402 (-16.78)	-5.6440 (-19.39)
LNFRNT	0.8122 (9.77)	0.6796 (6.62)	0.6723 (5.09)	0.8467 (7.44)
LNFRATE	0.8851 (14.88)	0.9427 (19.94)	0.8102 (15.94)	0.7053 (15.52)
CASH	-0.0323 (-0.78)*	-0.0511 (-1.01)*	-0.2226 (-4.76)	-0.0458 (-0.89)*
EFFTAX	0.0092 (4.63)	-0.0215 (4.91)	-0.0108 (4.23)	-0.0096 (4.34)
		r=10%		
CONSTANT	-5.0569 (-17.87)	-4.6330 (-17.36)	-4.1402 (-12.24)	-4.1099 (-14.12)
LNFRNT	0.8122 (9.77)	0.6796 (6.62)	0.6723 (5.09)	0.8466 (7.44)
LNFRATE	0.8851 (14.88)	0.9427 (19.94)	0.8102 (15.94)	0.7053 (15.52)
CASH	-0.0323 (-0.78)*	-0.0511 (-1.01)*	-0.2226 (-4.76)	-0.0458* (-0.89)
EFFTAX	-0.0426 (4.63)	-0.0998 (4.91)	-0.0503 (4.23)	-0.0443 (4.34)
R^2	0.7083	0.7231	0.7368	0.7093
# of OBS	140	168	115	119
F-RATIO	F(5,135)-9275.5	F(5,163)-8211.9	F(5,110)-9059.3	F(5,114)-11239.5
e'e	6.18	10.70	4.73	4.68

NOTES: The figures in parentheses beneath the regression coefficients are the estimated t-ratios; an asterisk outside the parenthesis indicates insignificance at conventional confidence levels.

The statistics reported under the regression coefficients pertain to the regressions using both r=2% and r=10%.

TABLE II

POOLED REGRESSION - r=2% and r=10%		
X	ESTIMATED COEFFICIENTS	
	r=2%	r=10%
CONSTANT	-4.6035 (-31.53)	-6.1377 (-42.04)
LNFRNT	0.7587 (14.42)	0.7587 (14.42)
LNFRATE	0.8429 (32.85)	0.8429 (32.85)
CASH	-0.0853 (-3.54)	-0.0853 (-3.54)
YQ^2	0.0219 (0.82)*	0.0219 (0.82)*
YQ^3	0.1033 (3.38)	0.1033 (3.38)
YQ^4	0.1783 (5.58)	0.1783 (5.58)
EFFTAX	-0.0573 (9.11)	-0.0123 (9.11)
	R^2 = 0.7194 F(8,535) = 21574.6 e'e = 28.1530	

TABLE III

POOLED REGRESSION - r=2% and r=10%

X	ESTIMATED COEFFICIENTS	
	r=2%	r=10%
CONSTANT	-10.1689 (-67.93)	-7.1005 (-47.44)
LNFRNT	0.7539 (14.47)	0.7539 (14.47)
LNFRATE	0.8434 (33.12)	0.8434 (33.12)
CASH (1,2,4)	-0.0433 (-1.59)*	-0.0433 (-1.59)*
CASH (3)	-0.2150 (-4.34)	-0.2150 (-4.34)
YQ^2	0.2466 (2.54)	0.2466 (2.54)
YQ^3	0.2489 (4.69)	0.2489 (4.69)
YQ^4	0.1958 (5.94)	0.1958 (5.94)
EFFTAX (1,3,4)	-0.0145 (7.52)	-0.0485 (7.52)
EFFTAX (2)	-0.0173 (6.01)	-0.0804 (6.01)
	R^2 = 0.7217 F(10,533) = 17654.8 e'e = 27.4222	

NOTES: The figures in parentheses beside CASH and EFFTAX refer to the quarters in which the estimated coefficients are constrained to be equal.

TABLE IV

ESTIMATED TAX COEFFICIENTS FOR VARIOUS EQUATIONS

VARIABLES		EQUATIONS				
Dependent	Independent	A	B	C	D	R^2
ln V	ln T	0.0370 (1.8419)*				0.60
ln V	T/V		-0.3593 (14.6965)			0.71
(ln V)'	T/V'			-0.3761 (22.41)		0.79
ln V & (ln V)'	(T/V).d & T/V, T/V'				-0.0068 (-0.6656)* -0.3856 (-27.6035)	0.74
Number of Observations:		543	543	543	1094	

NOTES:

i) d=1 if actual observation, and 0 if generated observation.

ii) The other variables included in equation A-D are those listed in the text.

iii) The table is set up similar to that in the Wales-Wiens article to facilitate comparison of the regression results on improved and unimproved properties.

APPENDIX

I. Testing the Equality of the Quarterly Slope Coefficients

The test for the equality of all the quarterly slope coefficients of a given variable is briefly outlined here. In the pooled or constrained regression there are 8 free coefficients, and in the quarterly regressions 20 free coefficients; thus 12 restrictions are imposed when the data is pooled. In comparing residual sum of squares (e'e), the following test statistic is formed:

$$\frac{e'e_c - e'e_{uc}}{e'e_{uc}} \; \frac{k}{R} \approx F_{kR} \qquad (1)$$

where k is the number of degrees of freedom, R is the number of restrictions, and $e'e_c$ and $e'e_{uc}$ are the residual sum of squares from the pooled regression and from the four quarterly regressions respectfully (the residuals from the four quarterly regressions are totalled).

Substituting values for k, R, $e'e_c$ and $e'e_c$ and e, $e'e_{uc}$ into (1) we get

$$\frac{(28.1530 - 26.29187)}{26.29187} \; \frac{523}{12} \approx F_{12,523}$$

$$3.08 \approx F_{12,523}$$

Referring to a statistical table we find that the critical values for $F_{12.523}$ at the 90 and 95 per cent confidence levels are 2.21 and 1.77 respectfully. Thus the hypothesis that the slope coefficients are equal is rejected.

II. Testing the Equality of Some of the Quarterly Slope Coefficients

Instead of constraining all the coefficients to be equal across all quarters, the coefficient of EFFTAX is allowed to vary in the second quarter, and that of CASH to vary in the third quarter - this regression is referred to as the constrained regression. We then test the hypothesis that the quarterly relations are identical to the constrained relation. In the constrained relation there are now 10 free coefficients and thus only 10 restrictions. Thus substituting new values for k, R, $e'e_c$ and $e'e_{uc}$ into (1) we get

$$\frac{(27.4222 - 27.20998)}{27,20998} \; \frac{523}{10} \approx F_{10,523}$$

$$.04 \approx F_{10,523}$$

The critical values of $F_{10,523}$ are 1.84 and 2.36 at 95 and 90 per cent confidence levels respectfully. Thus the hypothesis that the specified coefficients are equal cannot be rejected.

III. Testing the Equality of the Two Constrained Regressions

In this test the constrained regression is the one in which all the quarterly coefficients are constrained to be equal, and the unconstrained regression the one in which EFFTAX and CASH are allowed to vary. There are 10 free coefficients in the unconstrained regression, 8 in the constrained relation and thus only 2 restrictions being imposed. Substituting new values for k, R, $e'e_c$, and $e'e_{uc}$ into (1) we get

$$\frac{(28.1530 - 27.422)}{27.422} \; \frac{533}{2} \approx F_{2,533}$$

$$7.10 \approx F_{2,533}$$

The critical values for $F_{2,533}$ are 2.61 and 3.83 at the 95 and 90 per cent confidence levels respectfully. Thus the hypothesis that these two relations are the same is rejected.

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