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Do Tax-Deferred Exchanges Impact Purchase Price? Evidence from the Phoenix Apartment Market

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Many authors have commented on the compliance risk associated with tax-deferred exchanges. However, no published studies explicitly address whether the risks associated with the exchange process impact the price at which exchanged assets trade. Using a unique data set that documents transactions for nondirect exchanges, this study examines the price impact of tax-deferred exchanges on apartment transactions in the Phoenix, Arizona, market. Consistent with the price pressure hypothesis originally developed by Scholes (1972) and Kraus and Stoll (1972) and the tax capitalization hypothesis proposed by Oates (1969), the data show that exchange participants pay an economically significant premium to acquire replacement assets. A conventional hedonic price index is generated to investigate the rational bounds of the exchange premium.

In properly functioning capital markets, assets trade at the discounted value of future cash flows. However, market frictions can cause deviations from the equilibrium market price. Consistent with the price pressure hypothesis developed by Scholes (1972) and Kraus and Stoll (1972) and the tax capitalization hypothesis proposed by Oates (1969), the constraints associated with a Section 1031 asset exchange may result in transactions at nonequilibrium prices. This paper advances the literature by estimating empirically the impact of Section 1031 exchanges on the transaction price of apartment properties.

A tax-deferred exchange, frequently referred to as a Section 1031 exchange, can enhance the investment value of real property by deferring the tax liability associated with the disposal of appreciated real estate. The original IRS code on tax-deferred exchanges of real property was issued in 1921. However, before *Starker v. United States* in 1979, tax-deferred exchanges had to be executed

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simultaneously. The simultaneity requirement created a major hindrance to the execution of exchanges because of the difficulty of synchronizing the close of escrow on two or more complex transactions. Hence, few transactions were involved in the tax-deferment offered by a Section 1031 exchange (Goodman 1980).

In the Starker case, the court held that nonsimultaneous exchanges qualify for tax-deferred status (see Colwell and Dehring 2001 for an analysis of simultaneous vs. nonsimultaneous exchanges). Unfortunately, the taxing authority originally gave little administrative guidance on the proper execution of nonsimultaneous exchanges. While case law developed in the 1980s tended toward increasing leniency in the acceptable time period during which the two "legs" of the exchange transaction could be completed, the continued uncertainty surrounding the allowable parameters of the nonsimultaneous exchange discouraged widespread use (Rier 1985).

The Tax Reform Act of 1986 and the IRS regulations issued in May 1991 reduced the uncertainty over the parameters of a qualifying exchange by clearly delineating the maximum time period over which both legs of the exchange could be completed. As a result of the clear specification of the legal requirements, usage of the 1031 exchange vehicle increased dramatically in the early 1990s (Frank 1995).

The requirements of a qualifying exchange are now well defined. Specifically, once an investor relinquishes title to a property, he or she must identify a replacement property within a 45-day period. The replacement property transaction must then be closed within a further 135 days. The time intervals constitute absolute deadlines even when the 45th or 180th day falls on a weekend or legal holiday (Frank 1995). Substantial compliance is not adequate to preserve the tax-deferred status of the exchange.

Identification of potential replacement candidates within the 45-day time limit is frequently a binding constraint (Cuff 1998a). Failure to properly identify (and document) the potential replacement property within the specified time frame nullifies the tax-deferred status of the sale. As Hudson (1998), Lynch (1998), and Raitz and Raitz (2000) note, the time pressure associated with the identification period limits the due diligence efforts of the participant seeking the replacement asset. In addition to the uncertainties of identifying an acceptable replacement property within the time limit, a plethora of obstacles (such as permits, appraisals, loan approvals, inspections, licenses, and the competence and willingness of the other party) are beyond the control of the replacement buyer and may impact his ability to close on the replacement asset within the 180-day limitation (Sommers 1988).

Many authors (*e.g.*, Rier 1985; Sommers 1988; Groebe 1989; Levine 1991; Frank 1995; Freedman 1995; Cuff 1997, 1998a, 1998b; Banoff, Lipton and Kanter 1998; Raitz and Raitz 2000; and Killip and DeLeo 2000) note the compliance risk associated with attempting an exchange, especially a nonsimultaneous exchange. In particular, if an investor relinquishes title to a property with substantial appreciation in anticipation of executing a tax-deferred exchange, the exchange participant may have compromised his or her bargaining position with other parties, including the sellers of potential replacement properties (Sommers 1988). While some of these obstacles can be avoided by conscientious identification of the replacement property before disposal of the relinquished property, anecdotal evidence suggests that a substantial proportion of sellers seeking to effect an exchange have not identified a replacement asset at the close of escrow on the relinquished property (Anonymous 1997).

Whereas the literature on tax-deferred exchanges is replete with articles that warn of the compliance risk associated with exchange transactions, we are aware of no studies that examine explicitly whether compliance risks impact the price paid for the replacement property.[1] If market participants are engaged solely in the economic decisions associated with the disposal and acquisition of the exchanged assets, then transactions involving exchange participants will be priced the same as nonexchange transactions. If, however, Section 1031 fundamentally alters the parameters surrounding the decision to enter into a transaction, then the price paid for properties involved in an exchange transaction may be affected. This article advances the literature by assessing whether exchange transactions create price differentials in the apartment market in a test city, namely Phoenix.

The paper is organized as follows: The next section develops the theoretical foundation and model used to examine potential price differentials associated with exchange transactions, the third section discusses the data used in this investigation, the next section presents the empirical results, and the last section summarizes findings and conclusions.

[1] In the context of nonexchange topics, Downs and Slade (1999) and Munneke and Slade (2000) include an exchange variable in hedonic price analysis of office properties. In both cases, the data did not allow for identification of the exchange status, for example, if the transaction was part of the exchange participant's relinquished property or the replacement property. It was only known that the transaction was part of an exchange. In both cases, the parameter on the exchange variable was positive and significant. Also, in a thoughtful and interesting paper, Colwell and Dehring (2001) analytically derive the parameters surrounding the choice between a simultaneous exchange and a nonsimultaneous exchange for farmland located on the urban periphery.

Theory and Model

Theoretical Foundation

Suppose an investor with a capital gains tax rate of τ_c has relinquished a property with taxable capital gains of γ dollars in anticipation of effecting an exchange. Without the exchange, the investor would incur an immediate tax liability of

$$\text{Tax} = \gamma\tau_c. \tag{1}$$

If the investor is successful in effecting the exchange, the tax liability will be deferred during the holding period of the acquired property. The benefit of the exchange process is that the present value of the tax liability is reduced, since payment of the tax is delayed. For an investor with an expected holding period for the acquired property of n years and a cost of capital of r, the present value of the deferred tax liability is

$$\text{Present value of deferred tax} = \frac{\gamma\tau_c}{(1+r)^n}. \tag{2}$$

Hence, the value of using the tax-deferment provisions of a Section 1031 exchange for an investor with an expected holding period of n years is

$$\text{Exchange benefit} = \gamma\tau_c - \frac{\gamma\tau_c}{(1+r)^n}. \tag{3}$$

In the absence of regulatory constraints, an investor seeking to acquire real property would be willing to pay, at most, an amount equal to the present value of the asset's cash flows discounted at the market required rate of return k, formally,

$$\text{Price} = \sum_{t=1}^{n} \frac{CF_t}{(1+k)^t}. \tag{4}$$

Two well-established theoretical constructs provide motivation for the existence of a price premium in transactions involving a tax-deferred exchange. First, Scholes (1972) and Kraus and Stoll (1972) suggest that the price of an asset can be affected by "temporary" changes in demand. Under their hypothesis, developed in the context of block trades and referred to in the literature as the Price Pressure Hypothesis (PPH), the imposition of a temporary increase in demand will result in trades above the equilibrium price described in Equation (4). Numerous authors document evidence of price and demand effects consistent with the PPH.[2] In particular, Harris and Gruel (1986) conclude that price increases associated with temporary demand changes are necessary in order to attract "passive suppliers of liquidity." Given the documented liquidity constraints or thin markets associated with real property (Moore 1987, Kluger and Miller 1990, and Hasbrouck 1991), the urgency imposed on exchange participants by the regulatory time constraints may result in price differentials consistent with the PPH.

Second, Oates (1969) suggests that tax capitalization may impact the price of a traded asset by asserting that increased tax liabilities will be capitalized into the value of the taxed asset, resulting in lower property values. Numerous authors document varying degrees of tax capitalization associated with property tax obligations.[3] In contrast, the sharing of tax benefits between buyer and seller of a replacement asset constitutes reverse tax capitalization. That is, the buyer forfeits some portion of his gains in order to induce the seller to provide the needed or desired asset in a timely fashion. In effect, the exchange participant is able to simply outbid buyers who are not simultaneously benefiting from the significant deferral advantage associated with a 1031 exchange.

To illustrate the potential impact of the PPH and the tax capitalization hypothesis, assume an investor has relinquished an appreciated asset in anticipation of completing a Section 1031 exchange. As the regulatory deadline for identification of replacement properties approaches, the would-be exchange participant must identify a replacement property or recognize the full amount of the gain from the sale of the relinquished asset. Faced with thin markets, the investor may be pressured to pay a premium for the replacement asset. From the reverse tax capitalization perspective, the investor may be willing to pass onto the seller some of the tax savings in order to entice the seller to relinquish the desired property. The exchange participant seeking a replacement property could pay a premium up to the value of the exchange benefit in Equation (3) and still be as well off as other market participants. That is, a rational would-be exchange participant who is confronted with a choice between recognition of a gain and deferment through an exchange could pay up to

$$\text{Max price} = \sum_{t=1}^{n} \frac{CF_t}{(1+k)^t} + \left[(\gamma \tau_c) - \frac{\gamma \tau_c}{(1+r)^n} \right]. \quad (5)$$

[2] See Dann, Mayers and Raab (1977), Mikkelson and Partch (1985), Harris and Gurel (1986), Shleifer (1986), Loderer, Cooney and Van Drunen (1991), Simon (1994a, 1994b), and Babbel *et al.* (2000).

[3] For more complete discussion of the tax capitalization hypothesis, see Oates (1973), Church (1994), Ihlanfeldt and Jackson (1982), Hendershott and Ling (1984), Yinger (1988), Do and Sirmans (1994), and Palmon and Smith (1998).

Stated as a percentage of the nonexchange market price in Equation (4), the exchange participant could pay a maximum rational premium of

$$\text{Max premium} = \frac{\gamma \tau_c - \dfrac{\gamma \tau_c}{(1+r)^n}}{\displaystyle\sum_{t=1}^{n} \frac{CF_t}{(1+k)^t}}. \tag{6}$$

If the exchange participant anticipates an infinite deferral of the tax liability (*e.g.*, through a series of deferrals and or strategic estate planning), the upper bound on the magnitude of the rational premium is

$$\text{Upper bound} = \frac{\gamma \tau_c}{\displaystyle\sum_{t=1}^{n} \frac{CF_t}{(1+k)^t}}. \tag{7}$$

To the extent that price pressure or reverse tax capitalization results in the acquisition of a replacement property for less than the maximum premium defined in Equation (6), the exchange participant will be better off even though the price paid is greater than the equilibrium price in Equation (4). Hence, the exchange participant may still be better off even though the exchange can result in disadvantageous price premiums.

Model

In order to determine if price premiums are paid in tax-deferred exchanges, we estimate the following single equation, reduced form price function, to explain the price of apartment properties:

$$\begin{aligned} \text{Ln}\,PRICESF = {} & \alpha_0 + \alpha_1 \text{Ln}\,UNITS + \alpha_2 COVERPARK + \alpha_3 AGE \\ & + \alpha_4 AGESQ + \alpha_5 POOL + \alpha_6 CLUB + \alpha_7 LAUNDRY \\ & + \alpha_8 TENNIS + \sum_{i=2}^{3} \beta_i CONDITION_i \\ & + \sum_{i=2}^{3} \delta_i EXCHANGE_i + \sum_{i=2}^{7} \phi_i GEOAREA_i \\ & + \sum_{i=2}^{10} \lambda_i TIME_i, \end{aligned} \tag{8}$$

where,

Ln *PRICESF* = the natural log of the sale price per square foot of the property.

Ln *UNITS* = the natural log of the number of units in the complex.

COVERPARK = the number of covered parking spaces.

AGE = age of building(s) in years.

AGESQ = age squared.

POOL = a binary variable for the presence of a swimming pool (= 1 if present).

CLUB = a binary variable for the presence of a clubhouse (= 1 if present).

LAUNDRY = a binary variable for the presence of a laundry facility (= 1 if present).

TENNIS = a binary variable for the presence of a tennis facility (= 1 if present).

CONDITION = condition of the property, based on inspection. The categories include better than average, average, worse than average. Each category is included in the structural model as a binary variable except average, which is suppressed.

EXCHANGE = exchange status of the transaction; the possibilities include a buyer exchange (the purchaser is acquiring the asset as a replacement property in a qualified Section 1031 exchange), a seller exchange (the seller is relinquishing the property as part of a Section 1031 exchange), and nonexchange (neither the purchaser nor seller is involved in a qualified exchange). Each category is included in the structural model as a binary variable except nonexchange, which is suppressed.

GEOAREA = geographic location of each transaction. The geographic areas include Central Phoenix, North Phoenix, West Phoenix, Northwest Valley, Scottsdale, Tempe, and East Valley. Each geographic area is included in the structural model as a binary variable except Central Phoenix, which is suppressed.

TIME = quarterly time periods from third quarter 1995 through fourth quarter 1997. Each quarterly time period is included in the structured model as a binary variable except third quarter 1995, which is suppressed.

The dependent variable is specified as the natural logarithm of price per square foot. This form is common in hedonic pricing literature and is consistent with the approach used by Linneman (1980). One of the advantages of this form is that it gives less weight to extremely high values than does an untransformed variable (de Leeuw 1993). The number of units variable appears in logarithmic form, thus allowing the coefficient to be interpreted as an elasticity. In this case, Ln *UNITS* is expected to be negative, suggesting that price per square foot will decrease at a declining rate with respect to the number of units in the complex. This specification is consistent with the result found by de Leeuw (1993) and implies that small economies of scale are generally expected with larger projects. In the Phoenix market, because the extreme summer heat results in a high preference for covered parking stalls, covered parking is likely to positively impact rents, and thus value. A logarithmic transformation of the parking variable is not possible because some complexes have no covered parking. The *AGE* of the complex is expected to be negatively related to price per square foot. Because properties typically depreciate at a slower rate over time, the *AGESQ* variable is included to capture the declining rate of depreciation.

The impact of valued amenities, such as those represented by the variables *POOL*, *CLUB*, *LAUNDRY*, and *TENNIS*, is expected to be positive. Each property in the sample was inspected and rated as to overall condition. The inspection process, performed by Costar Group, Inc. (formerly Comps InfoSystems, Inc.), resulted in each property being assigned to one of five condition categories: excellent, good, average, fair, or poor. In the data set used in the empirical investigation, relatively few properties were categorized as excellent or poor; therefore, excellent and good were combined to create a new variable, better than average. Also, fair and poor were combined to create a new variable, worse than average. In our operational model, the suppressed category is "average." The parameter estimated for *BETTER THAN AVERAGE* is expected to be positive, while the parameter for *WORSE THAN AVERAGE* is expected to be negative.

The true variables of interest are the exchange variables, *BUYER EXCHANGE* and *SELLER EXCHANGE*. The PPH, as well as the reverse tax capitalization hypothesis, suggests that the *BUYER EXCHANGE* variable will have a positive parameter. Indeed, if the parameter on the buyer exchange variable is significant and positive, we will conclude that the regulation is impacting the price of the replacement property.

The intuition supporting the impact of the *SELLER EXCHANGE* variable is more obtuse. Under the PPH, sellers are unlikely to accept a below-market price in order to rush into the risks associated with finding and closing on the replacement property. Given that no incentive to rush into the sale of the relinquished property exists (indeed, the incentive is to delay closing), the seller of a relinquished property experiences no regulation-induced urgency.[4] Under this reasoning, the coefficient on the *SELLER EXCHANGE* variable should be insignificant. However, the sale of the relinquished asset could be motivated by a desire to buy some other property that the firm finds attractive. Consistent with the tax capitalization hypothesis, the seller may be willing to share some portion of the expected tax benefits of the deferred exchange with the purchaser of the relinquished asset under these circumstances. With this reasoning, a negative coefficient on the *SELLER EXCHANGE* variable would be expected. Thus, we expect that the coefficient on the *SELLER EXCHANGE* variable will be either insignificant or negative.

The Phoenix metropolitan area consists of numerous contiguous cities and economic submarkets. To control for the differences in location that may impact price, geographic dummy variables are incorporated into the model. The data provider has segregated the metropolitan area into various submarkets that are commonly referred to by real estate practitioners in this market. These submarket classifications have been used to construct the location variables. Quarterly time dummy variables are also incorporated into the model to capture any intertemporal price changes that have occurred during the period under investigation.

In any empirical analysis without observable determinants, the possibility of omitted variable bias exists. For omitted variables to distort our findings on the price impact of exchanges, the omitted variable would have to have explanatory power, be correlated with the exchange variable, and not be explained by the included set of independent variables in Equation (8). Viewing our inference structure as a standard omitted variable test for the impact of an exchange transaction (*e.g.*, Holmes and Horvitz 1994, Hunter and Walker 1996, and Phillips-Patrick and Rossi 1996), the potential impact of omitted hedonic variables is muted. While omission of a hedonic variable such as story height may impact other hedonic variables, such as parking, it is more difficult to see how omission of a hedonic variable would significantly impact the coefficient of nonhedonic variables, such as the variables of interest, namely the exchange variables.

[4] Also relevant is the possibility of buying the replacement property first and then marketing the property to be relinquished in a process known as a *reverse Starker exchange*. The advantage of the reverse Starker is that, because Section 1031 is stated in the negative and is forward looking, the stringent time constraints that apply to a normal exchange do not apply to the reverse exchange. Several authors, including Sommers (1988), Killip and DeLeo (2000), and Raitz and Raitz (2000), propose reverse exchanges as a way to circumvent the compliance risk associated with forward exchanges. However, the administrative costs associated with a reverse Starker exchange are significantly higher. The data set used in this study contains no transactions involved in a reverse Starker exchange.

Data

The primary data used for this study include apartment transactions from the Phoenix metropolitan area from September 1995 through December 1997. The aggregate data set includes 784 transactions and was provided by CoStar Group, Inc.[5] CoStar investigates all apartment transactions exceeding $150,000 by physically inspecting each property and confirming the particulars of the transaction with the relevant parties, including buyer, seller, and broker. This includes verification of the exchange status of the transaction. During the inspection process, the inspector documents the physical characteristics of the property and provides a subjective estimate of the condition of the property. Transactions data prior to September 1995 were obtained, but details pertaining to the exchange status of each property were not documented prior to this time. Therefore, the sample is limited to the period from September 1995 through December 1997.

Apartment markets are often categorized as residential (two to four family units) or commercial multifamily (five or more units). The original data set contained 80 transactions involving properties with four or fewer units. The analysis reported here focuses only on the multifamily transactions. In addition, 34 other transactions were eliminated because of missing hedonic variables or obvious data input errors. Three transactions were found to be part of a direct exchange; 11 transactions were found to be part of both a seller's exchange and a buyer's exchange. Because of the paucity of direct exchange transactions, these were eliminated from the study.[6] The 11 transactions found to be part of both sides of an exchange were also deleted so as to eliminate possible bias that these unusual transactions might pose to the model. The resulting data set consists of 656 transactions. Because of the level of detail pertaining to these transactions, particularly the information relating to the exchange status, the data set provides a unique opportunity to examine the impact of exchange status on sales price. Table 1 provides descriptive statistics of the data set.

The range in sales price and building area illustrate a large variation in the value and size of the properties. In addition, a mean building age of 24 years combined with a standard deviation of 12 years suggests that many of the apartments in the Phoenix area are relatively newer, especially compared with apartment properties located in northeastern cities. As is expected in the Phoenix area, over 50% of the properties have a swimming pool. Approximately 15% have an onsite clubhouse facility. A majority of the properties (79%) is considered in average

[5] CoStar Group, Inc. (formerly Comps InfoSytems, Inc.) investigates and compiles real estate transaction data in many cities in the United States, including Phoenix, Arizona. Summaries of the transactions are provided to interested parties on a subscription basis. We thank Craig Farrington and Dan Prevo for their generous assistance with the data.

[6] Colwell and Dehring (2001) examine direct exchanges involving farmland. Investigation of apartment properties in Phoenix found very few direct exchanges.

Table 1 ■ Descriptive statistics of the transaction data, apartment properties, Phoenix, September 1995–December 1997.

Variable	Mean	St. Dev.	Minimum	Maximum
Sales price	$2,822,000	$4,812,000	$150,000	$33,600,000
Number of units	82	109	5	762
Building area (sq. ft.)	63,762	86,547	3,869	553,000
Building age	24	12	0	66
Covered parking spaces	62	110	0	841

Frequency of Dichotomous Variables

Variable	Mean	Observations
Total observations		656
Project amenities		
Swimming pool(s)	0.5793	380
Clubhouse	0.1555	102
Laundry	0.3445	226
Tennis court(s)	0.0915	60
Condition		
Better than average	0.0732	48
Average	0.7911	519
Worse than average	0.1356	89
Exchange variables		
Non-exchange	0.8796	577
Buyer exchange	0.1082	71
Seller exchange	0.0122	8
Geographic Areas		
Central Phoenix	0.4284	281
North Phoenix	0.1402	92
West Phoenix	0.0427	28
Northwest Valley	0.0777	51
Scottsdale	0.0686	45
Tempe	0.0777	51
East Valley	0.1646	108
Quarterly Time Periods		
3rd quarter 1995	0.0305	20
4th quarter 1995	0.0777	51
1st quarter 1996	0.1113	73
2nd quarter 1996	0.1098	72
3rd quarter 1996	0.1326	87
4th quarter 1996	0.1159	76
1st quarter 1997	0.1006	66
2nd quarter 1997	0.0961	63
3rd quarter 1997	0.1113	73
4th quarter 1997	0.1143	75

Notes: The data include transactions of apartment properties from September 1995 through December 1997. All properties exceed four units and are located within the Phoenix metropolitan area. The data were obtained from the CoStar Group, Inc. (formerly Comps InfoSystems, Inc.). Employees at the CoStar Group physically inspect each property and confirm transaction details with relevant parties. The data are then provided to interested parties on a subscription basis.

condition, while approximately 7% are better than average and 14% worse than average. The exchange variables, which are the focus of this study, include buyer's exchange, seller's exchange, and nonexchange. Approximately 11% of the transactions were part of a buyer's exchange, while only 1% were part of a seller's exchange. The remaining 88% were not involved in an exchange. The paucity of seller exchange transactions limits the reliability of any results associated with this variable.

Using geographic sectors provided by the data providers, the Phoenix metro area is segregated into seven geographic market areas: Central Phoenix, North Phoenix, West Phoenix, Northwest Valley, Scottsdale, Tempe, and East Valley. The Central Phoenix market dominates the sample, with over 42% of all apartment transactions. East Valley is the second largest area, with just over 16% of the transactions. The remaining areas range from 4 to 14% of the transactions. Including location variables in the model allows for proper control of possible spatial characteristics that may influence the sales price.

The period under investigation (September 1995 through December 1997) includes 10 quarters of data. Other than the first 2 quarters, which have fewer transactions, the number of transactions per quarter is relatively stable. Segregating the transactions by quarter allows for proper control of intertemporal price changes that may impact the hedonic pricing model.

One important caveat regarding the data set is pertinent. Many attribute-based hedonic valuation models include proxies for operating income or expenses and unit-mix. Unfortunately, complete income and expense data are unavailable. We believe that the combination of hedonic variables included, such as *AGE*, *CONDITION*, and *LOCATION*, significantly homogenize differences in income. Additionally, while total number of units is included in the analysis, the details of unit-mix (1 bedroom, 2 bedroom, 3 bedroom) are not included due to data constraints.[7] For the omission of a proxy to bias our results, differences

[7] The available data include unit-mix information for about 80% of the observations. In order to assess the potential bias from the omission of unit-mix controls in the full sample results reported in Table 2, we reestimated our model using the subset of observations for which we have unit-mix data. Four points are relevant. First, the unit-mix variables (1 bedroom units, 2 bedroom units, 3 bedroom units) are insignificant in explaining the price per square foot of the complex. Second, the addition of the unit-mix control variables increases the adjusted R^2 by only 0.013 (from 0.5047) in the relevant subset. Third, no significant correlation exists between the exchange variable and the unit-mix variables for those observations for which we have unit-mix data. Finally, there is no qualitative change in the variable of interest (*BUYER EXCHANGE*) when our model is run on the subset of data that includes unit-mix controls. From the analysis of the subset, bias stemming from the omission of unit-mix controls appears unlikely; apparently the existing model captures the explanatory information that unit-mix variables would generally provide.

between properties that are not explained by the included variables would have to be highly correlated with the exchange variables. While this seems unlikely, the reader is cautioned to consider this caveat when interpreting the empirical results.

Empirical Results

The objective of our model is to determine if price differentials occur in transactions involved in a tax-deferred exchange. Table 2 shows the estimation results

Table 2 ■ Regression results.

Explanatory variables	Parameter	t-Statistics
Intercept	3.766	(44.8)
Natural log of units	−0.0586*	(4.1)
Age	−0.0215*	(6.4)
Age squared	0.0004*	(6.3)
Covered parking	0.0008*	(6.1)
Swimming pool	0.0907*	(3.8)
Clubhouse	0.1009*	(2.9)
Laundry	0.0007	(0.03)
Tennis court	0.0234	(0.6)
Better than Avg. condition	0.1723*	(4.0)
Worse than Avg. condition	−0.1924*	(6.6)
Buyer exchange	0.0763*	(2.5)
Seller exchange	−0.0119	(0.1)
North Phoenix	−0.0136	(0.5)
West Phoenix	−0.1479*	(3.1)
Northwest Valley	−0.0432	(1.2)
Scottsdale	0.2895*	(7.3)
Tempe	0.2568*	(7.0)
East Valley	0.1438*	(5.1)
4th quarter 1995	−0.0465	(0.7)
1st quarter 1996	0.0252	(0.4)
2nd quarter 1996	0.0375	(0.6)
3rd quarter 1996	0.0589	(0.9)
4th quarter 1996	0.0206	(0.3)
1st quarter 1997	0.0600	(0.9)
2nd quarter 1997	0.1832*	(2.9)
3rd quarter 1997	0.1429*	(2.3)
4th quarter 1997	0.2039*	(3.3)
Adjusted R-square	0.5047	

Dependent variable is natural log of sales price per square foot of building area. The suppressed dummy variables include average condition, nonexchange, Central Phoenix, and 3rd quarter 1995. The absolute values of the t-statistics are presented in parentheses.

*Significant at the 0.05 level.

of Equation (8). Of the physical characteristic variables, all but *LAUNDRY* and *TENNIS* exhibit the expected sign and are significant at the 0.05 level. The coefficient on the Ln*UNITS* variable suggests that price per square foot declines at a decreasing rate as the number of units in the complex increases. The coefficients on the *AGE* and *AGESQ* variables confirm that price per square foot also declines at a decreasing rate with respect to property age. The remaining project characteristic variables, such as *COVERPARK*, *POOL*, and *CLUB*, are positive and significant, as expected. With regard to the condition variables, coefficients conform to the intended classification system and are statistically significant. Table 2 shows a positive and significant coefficient on *BETTER THAN AVERAGE* and a negative and significant coefficient on *WORSE THAN AVERAGE*, as predicted. Four of the six geographic market variables are significantly different from those of Central Phoenix, suggesting that geographic location is important in explaining apartment prices in the Phoenix area.

Three of the four 1997 quarterly time dummy variables are positive and significant, indicating that prices are increasing over the study period. Given that the dependent variable is price per square foot, the adjusted R-square of 0.50 suggests that the model does a good job of explaining the sales price per square foot for apartments in the Phoenix market.[8]

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[8] To ensure the reliability of the regression results, we examined the error terms. Specifically, we conducted a runs test derived by Geary (1970) to determine if the error terms were independent and normally distributed. The test confirmed that the results satisfy the classical assumptions of independence and normality. In addition to the runs test, we tested for spatial autocorrelation by calculating the test statistic "Moran's I," defined as

$$(N/S)\left(\sum_{i=1}^{N}\sum_{j=1}^{N} W_{ij} Z_i Z_j\right) \Big/ \left(\sum_{i=1}^{N}\sum_{j=1}^{N} Z_i^2\right),$$

where N is the number of observations, S is the sum of the weight matrix, and W is the weight matrix. Z is the random variable minus the mean evaluated for regions i and j (see Oland 1987). We followed the approach that is standard for a linear regression model, in which the residuals from and OLS regression are input as the Zs in the model (see Anselin 1988). In addition, we have calculated Moran's I for a class of lag distances. This implies that all the elements in the weight matrix are initially set to 1 and then weighted by the distance between pairs of points divided by the class of lag distances. Specifically, we have used the standard of 80% of the greatest distance between two points as the distance class. This distance is then broken into 10 intervals for which Moran's I is calculated. Thus, rather than measuring simple spatial contiguity, we have measured spatial correlation by distance. This results in greater robustness in our results (see Robertson 2000). The results can be graphed in a correlogram, in which the ten Moran's I scores are measured on the vertical axis and the distance intervals are measured on the horizontal axis. This process was first done for a regression model that contained no dummy variables for geographic location. Without the location dummy variables, the model suffered from minor spatial autocorrelation. Second, we performed the analysis for the residuals from the regression that did contain geographic dummy variables. By comparing the results from the two analyses, it is apparent that the location dummy variables successfully corrected the spatial autocorrelation.

The parameter on *SELLER EXCHANGE* is insignificant. As noted above, the PPH suggests that sellers would not be subject to the same regulation-induced urgency as buyers. Hence, the insignificant coefficient on the *SELLER EXCHANGE* variable is not surprising. We do caution the reader that the paucity of *SELLER EXCHANGE* observations may be influencing the coefficient; therefore, further investigation of this variable is warranted.

As conjectured (but never tested) by many articles on replacement property acquisition, the coefficient on *BUYER EXCHANGE* is positive and significant at the 0.05 level, statistical evidence that buyers pay a premium for replacement assets in a tax-deferred exchange. The positive impact on the purchase price of the replacement asset is consistent with the PPH as it pertains to the time constraints and thin markets associated with tax-deferred exchanges in real estate markets. The result is also consistent with the reverse tax capitalization hypothesis.

Given the statistical significance of the *BUYER EXCHANGE* variable, interpretation of economic significance and rationality is pertinent. The estimated parameter on *BUYER EXCHANGE* is 0.0763, indicating that, *ceteris paribus*, an exchange participant acquiring a replacement asset pays a premium of approximately 7.9%.[9] A purchase price premium of this magnitude is economically meaningful to most observers. Consistent with the PPH and the tax capitalization hypothesis, the data support the hypothesis that a buyer acquiring a property to complete a nonsimultaneous Section 1031 exchange pays a significant premium to outbid other potential purchasers in order to complete the transaction within the required time frame.

The recent history of the Phoenix market is pertinent to the interpretation of the rationality of the exchange impact. Real estate prices in Phoenix experienced large declines in the late 1980s and early 1990s. Our data, which come from the 1995 to 1997 period, represent transactions that occurred after a strong recovery in prices. Hence, our sample may represent exchanges with larger than average capital gains. Since the benefits of an exchange increase in proportion to the magnitude of the gains to be deferred, the magnitude of the parameter on *BUYER EXCHANGE* may be greater than in other settings.

The economic rationality of the estimated coefficient for *BUYER EXCHANGE* can be formally assessed using Equation (7), the rational upper bound on the magnitude of the premium for the purchase of the replacement asset. Since

[9] The coefficient on the *BUYER EXCHANGE* variable can be transformed into an indication of the percentage of price increase by using the relationship *PERCENT INCREASE* $= 100[e^{0.0763} - 1]$ or 7.9% (Halvorsen and Palmquist 1980).

the upper bound assumes infinite deferment, the rational limits of the exchange premium are an increasing function of the amount of the gain to be deferred and the tax rate on capital gains. However, the tax rate on capital gains during our sample period is constant at 28% (*The 1999 U.S. Master Tax Guide* 1998).[10] Given that the tax rate is constant, the rational upper bound is reduced to a function of the amount of the capital gain to be deferred and the nonexchange value of the replacement property.

Unfortunately, the amount of the capital gain to be deferred in the exchange transactions, γ, is not available. However, examination of local market changes can provide insight into the potential for capital gains. Hence, we estimate a conventional hedonic price index for the Phoenix apartment market from 1990 through 1997.[11] Following Clapp, Giaccotto and Tirtiroglu (1991), as well as Knight, Dombrow and Sirmans (1995), we included all variables shown in Equation (8) as well as a vector of annual time dummy variables (1990–1997) and generated the regression results. The coefficients on the time variables capture the intertemporal pure price change. The index is constructed by taking the antilogarithm of the time coefficients and normalizing to unity. Figure 1 illustrates the price index for the Phoenix apartment market from 1990 through 1997.[12]

The index shows the market in decline from 1990 through 1992. However, after the trough, the market experiences a strong recovery through the end of the study period. In fact, the market experiences a 63% increase in prices between 1992 and 1997. Given that the value of an exchange increases with the amount of the capital gain to be deferred, we expect the average appreciation for relinquished properties sold in anticipation of an exchange to be greater than the appreciation for properties generally. Overall, the conventional hedonic index shows that the Phoenix apartment market experienced large increases in price from 1992 through 1997, suggesting a strong real estate market and the possibility of large capital gains that could be sheltered with a 1031 tax-deferred exchange.

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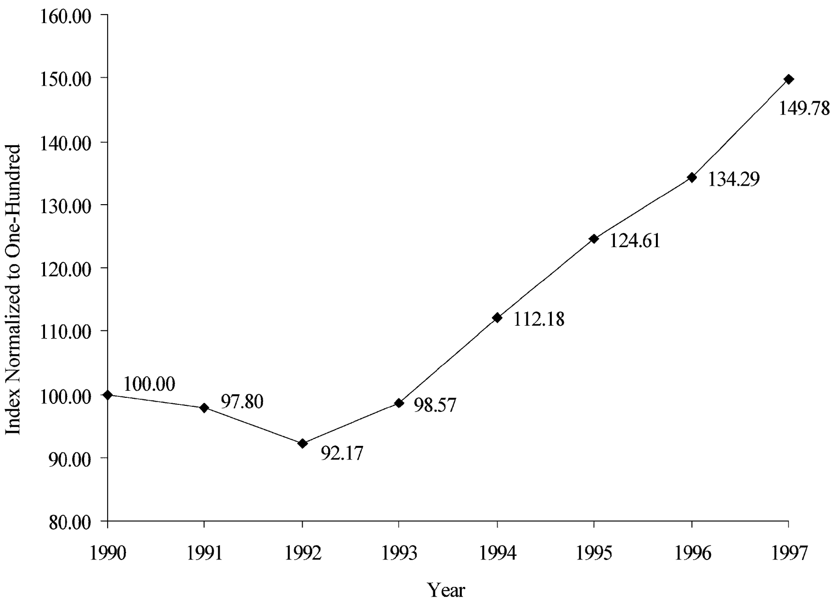
[10] The capital gains tax rate changed from 28% to 20% in May of 1997 (the last year of our sample). The rate at which capital gains were taxed in 1997 depended on when the relinquished asset was sold. Hence, if the full 180-day period was consumed in the exchange process, only observations from November and December of 1997 would involve replacement asset acquisitions in which the gains on the sale of the relinquished asset would have been subject to the new capital gains rate. Therefore, the 28% capital gains rate is considered more pertinent in this analysis.

[11] We do not have detailed information on the exchange status of transactions prior to September 1995. Therefore, in an effort to negate any adverse bias, we have eliminated all observations involved in an exchange from the hedonic price index estimation.

[12] See Table 3 in the Appendix for descriptive statistics of the entire data set.

Figure 1 ■ Hedonic price index of Phoenix apartment properties, 1990–1997.



Given the lack of data on the amount of capital gains to be deferred in each buyer exchange transaction, we are unable to estimate precisely the rational upper bound for the exchange premium. However, we can determine the relative magnitude of γ that places our estimated premium of 7.9% within the rational bounds. To determine whether our estimated exchange premium is rational, we substitute Equation (4) for the nonexchange value of the replacement asset into the denominator of Equation (7), the theoretical value of the rational upper bound, and set the resulting equation equal to the estimated premium of 7.9%, as

$$\frac{\gamma \tau_c}{PRICE} = 0.079. \tag{9}$$

Recalling that τ_c is constant at 28% during our sample period and solving Equation (9) for γ yields

$$\gamma = 0.2821(PRICE). \tag{10}$$

Hence, Equation (10) shows that the estimated exchange premium of 7.9% is rational as long as γ, the amount of capital gain to be deferred, is greater than or equal to 28.21% of the purchase price of the replacement asset. Given the 63% average price appreciation identified by the hedonic price index (and the tendency of exchanged assets to have greater than average capital gains), a rationality requirement that capital gains exceed 28% of the purchase price of the replacement asset does not seem unreasonable. While available data do not allow direct estimation of the rational upper bound, we conclude that the estimated premium is both economically meaningful and reasonable.

Summary and Conclusion

Although codified in the 1920s, the tax deferment provisions of a Section 1031 exchange experienced very little use in real property transactions for a number of decades because of uncertainty surrounding the parameters of a qualifying exchange. However, much of the uncertainty was resolved in 1991 when the IRS issued final guidelines delineating the precise time constraints for the purchase of a replacement asset in a nonsimultaneous exchange. Under these guidelines, a nonsimultaneous exchange qualifies for tax-deferred status under Section 1031 if (i) the replacement property is identified within 45 days of the close of the relinquished property, and (ii) the purchase of the replacement property is completed within 180 days of the close of the relinquished asset. The increased certainty of the criteria for a qualifying exchange led to widespread use of tax-deferred exchanges in the 1990s.

While the clear specification of the time requirements for an exchange eliminated uncertainty of the legal requirements for tax deferment, the relative shortness of the allowable time interval between the sale of the relinquished property and the completion of the exchange creates significant compliance risk. Many authors warn of the compliance perils associated with attempting nonsimultaneous exchanges (*e.g.*, Rier 1985; Sommers 1988; Groebe 1989; Levine 1991; Frank 1995; Freedman 1995; Cuff 1997, 1998a, 1998b; Banoff, Lipton and Kanter 1998; Raitz and Raitz 2000; and Killip and DeLeo 2000). Specifically, the thin markets often associated with real property may be exacerbated by the severe time constraints associated with the exchange process to produce disadvantageous price premiums.

Consistent with the price pressure hypothesis and the tax capitalization hypothesis, economic intuition affirms that the combination of rigid time constraints, impaired negotiating position, and thin real estate markets could have an impact on purchase price for assets involved in an exchange. However, while many authors discuss the risks associated with the exchange process, no effort has been devoted to discerning empirically whether compliance risk impacts the purchase price of exchanged assets.

Using a unique data set, we estimate a hedonic pricing model that explains sales price for 656 apartment transactions in the Phoenix market. To the hedonic model, we add binary variables to indicate whether the transaction is part of an exchange. The results are convincing. While sales price of relinquished assets are not significantly impacted, the data clearly show that exchange participants pay a premium for replacement assets consistent with the price pressure hypothesis developed by Scholes (1972) and Kraus and Stoll (1972) and tax capitalization hypothesis by Oates (1969). The estimated coefficient on the *BUYER EXCHANGE* variable of 0.076 (p-value of 0.012) suggests that exchange participants pay an economically significant premium to acquire replacement assets.

The findings show that the regulatory constraints imposed by the requirements of a Section 1031 exchange materially alter the distribution of resources. The primary limitation of our results is the scope of the investigation. We examined one property type (apartments), in one market (Phoenix), over one time period (1995–1997). Given the magnitude of the impact of exchanges on sales price in this sample, there is a clear need to extend the research initiated here to other property types, regions, and time periods.

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Appendix

Table A1 presents descriptive statistics of the transaction data for apartment properties in Phoenix from 1990 to 1997.

Table A1 ■ Descriptive statistics of the transaction data, apartment properties, Phoenix 1990–1997.

Variable	Mean	St. Dev.	Minimum	Maximum
Sales price	$2,023,300	$3,898,300	$80,000	$33,600,000
Number of units	75	110	5	1,140
Building area (sq. ft.)	57,492	85,455	1,813	779,390
Building age	21	13	0	81
Covered parking spaces	55	108	0	1,140

Frequency of Dichotomous Variables

Variable	Mean	Observations
Total observations		2,351
Project amenities		
Swimming pool(s)	0.5104	1,200
Clubhouse	0.1285	302
Laundry	0.1969	463
Tennis court(s)	0.0629	148
Condition		
Better than average	0.2063	485
Average	0.6640	1,561
Worse than average	0.1297	305
Geographic areas		
Central phoenix	0.3866	909
North phoenix	0.1821	428
West phoenix	0.0357	84
Northwest valley	0.0770	181
Scottsdale	0.0774	182
Tempe	0.0761	179
East valley	0.1650	388
Yearly time periods		
1990	0.0957	225
1991	0.1238	291
1992	0.1540	362
1993	0.1489	350
1994	0.1455	342
1995	0.1174	276
1996	0.1195	281
1997	0.0953	224

Notes: The data include transactions of apartment properties from 1990 through 1997. All properties exceed four units and are located within the Phoenix metropolitan area. The data were obtained from the CoStar Group, Inc. (formerly Comps InfoSystems, Inc.). Employees at the CoStar Group physically inspect each property and confirm transaction details with relevant parties. The data are then provided to interested parties on a subscription basis.