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ISSUES IN PUBLIC UTILITY VALUATION AND TAXATION

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SECURITY MARKET VALUES AND TAX-ADJUSTED FUNDAMENTAL VALUES FOR PUBLIC UTILITY PROPERTY

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Research improving techniques for constructing unbiased estimates of economic values for public utility property is justifiable from several perspectives. Regulators may properly adduce the rate of return on capital required by financial markets, yet if the net assets for computing the allowable return is misvalued, then the target earnings will be incorrect. Alternatively, changes in the economic value of net assets constitutes an estimate of economic depreciation. Regulators relying upon erroneous economic depreciation estimates will misstate recoverable costs. Indeed, a consistent and equitable approach to *ad valorem* taxation requires the unbiased estimation of economic values.

There are two generally accepted classes of accounting valuation models for fixed public utility property. Original historical cost (*HC*) models specify the association between the asset's acquisition cost and the decline throughout the asset's service life in productive capacity. It usually is recognized, however, that *HC* measurements are biased estimates of economic values because they ignore wealth revalutions resulting from changing price levels.

Current cost (*CC*) models incorporate effects of specific price changes. The preference for valuation estimates generated by *CC* models is summarized in the following specification:

$$MV_t = CC_t, \tag{1}$$

where MV_t is the market value of a firm or industry's liabilities at time *t* and CC_t is the current cost of its assets. The above specification is consistent with Tobin's [1969] argument that at equilibrium the *Q*-ratio (i.e., *MC/CC*) equals unity.

This study hypothesizes that the current cost model is systematically biased, i.e.,

$$MV_t = CC_t \times B_t, \tag{2}$$

where *B* is the systematic bias inherent with the current cost methodology. A model for measuring the bias is specified below. This alternative specification of economic value, *CC*x*B*, is referred to as the *tax-adjusted fundamental value* model (*FV*).[1]

An empirical test on the significance of the bias is obtained by econometrically estimating equation 2 in the following format,

$$\log MV_t = a_o + a_1 \log CC_t + a_2 \log B_t + e_t, \tag{3}$$

where e_t is a random error term and the *a*'s are estimated coefficients. According to the *CC* model depicted in equation 1, a_1=1 and a_o=a_2=0. According to the *FV* model depicted in equation 2, however, a_1=a_2=1 and a_o=0. In effect, the significance level for a_2 is an acid test on the incremental information content of the *FV* measurements.

This study proceeds as follows. First, the similarities and dissimilarities between the *CC* and *FV* valuation models are discussed. Next, measurements of *MV*, *CC*, and *B* are constructed for 3 utility industries over the 1968-

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85 annual sample period. Finally, the data are used to estimate the parameters in equation 3. A brief summary concludes the study.

Specification of Models for Valuing Fixed Assets

Generally, estimates of net asset values are obtained through reliance on a valuation model specifying the relation between the value of a used asset, the price of a new asset, and the decline throughout the asset's service life of potential productive capacity. Let $p_t q_t$ denote the gross investment in new fixed assets at time *t*. The price or nominal component, p_t, represents the specific price at time *t* of a new asset possessing one unit of productive capacity. The quantity or real component, q_t, represents the number of units of productive capacity acquired at time *t*. Hence, *pq* represents the fixed asset expenditure. Let d_j, for j=0,...,∞, represent the capacity depreciation schedule; d_j denotes the proportional decline in productive capacity occurring at the end of the *j*'th year of service. For example, with straight-line capacity depreciation over a four year service life $d_j = 1/4$ for $j=1,...,4$ and d_j=0 otherwise. Accumulated capacity depreciation for an asset concluding its *t*'th year of service equals $\Sigma^{t}_{j=o} d_j$.

The current cost (*CC*) valuation model partitions the time *t* fixed asset expenditure into its quantity and price components. The quantity component is net of accumulated capacity depreciation and the price component is set equal to the current specific price of a new asset. Accordingly, the current cost valuation measurement at time *s* for all net fixed assets is

$$CC_s \;=\; \sum_{t=o}^{\infty} p_s q_{s-t} \{1-\Sigma^{t}_{j=o} d_j\}. \tag{4}$$

This formulation of *CC* as a new price component less a depreciation component was adopted by over 90 percent of the preparers of *CC* data in the FASB [1979] Statement No. 33 supplementary disclosures (see Shriver [1987]). The traditional historical cost model (*HC*) for valuing assets is a special case of equation 4 in which p_{s-t} replaces p_s. When prices are constant, of course, the *HC* and *CC* models yield identical measurements.

An asset's tax-adjusted fundamental value (*FV*) is defined as the discounted sum of its expected after-tax cash flows. A zero net present value investment equilibrium in the fixed asset market is assumed so that for new assets *FV* and *CC* are identical. Downs [1992] shows that for the total stock of net fixed assets

$$FV_s = CC_s \times B_s, \tag{5}$$

where B_s is a specification for the theoretical bias inherent with the *CC* measurement,

$$B_s = (1-v-\tau Z_s)(1-H_s)/(1-D_s) + \tau Y_s. \tag{6}$$

The variables D_s and H_s are measures of discounted capacity depreciation for new assets and for the total asset stock, respectively. Z_s and Y_s are the discounted tax depreciation deductions per dollar of new assets and per dollar of total net assets, respectively, τ is the statutory income tax rate, and v is the rate of the investment tax credit. Explicit specifications for these variables and data sources are presented in Downs and Shriver [1993].

The general relation between tax-adjusted fundamental value and current cost is depicted in equatio 5. Under some conditions the bias *B* is unity, implying the *FV* and *CC* models yield identical valuation measurements. Differences between *FV* and *CC* arise, however, due to discounting and differential tax effects. In short, though, the *CC* model assumes that discounted after-tax cash flow is a constant proportion of current productive capacity whereas the *FV* model allows for differences between new and used assets in effective tax rates and in the ratio of current-to-future productive capacities.

Measurement of MV, CC, and B

To test whether the current cost model systematically misvalues public utility property, measurements for *MV*, *CC*, and *B* are collected over the 1968-85 annual sample period for three industries: Telephone Communications (SIC 481), Electric Services (SIC 491), and Natural Gas Transmission and Distribution (SIC 492).

The *MV* variable reported in table 1 represents the total market value of the industry's financial liabilities and is set equal to the sum of the market value of equity, short-term debt, and long-term debt. The industry market value of equity equals the sum across firms in each respective industry on *Compustat* of their number of common shares outstanding times year-end price per share (preferred equity is ignored). The industry market value of debt equals the sum across firms of the values of short-term and long-term debt. The procedure for valuing debt is adopted from VonFurstenberg, Malkiel, and Watson [1980] and sets the value of short-term debt equal to the book value collected for the industry from *Compustat*. Long-term debt is converted from book to market value by multiplying with price to book ratios. These ratios vary by year and by industry and rely on sampling annual bond price quotations from *Moody's Bond Record*.

The *CC* variable, also reported in table 1, represents the total current cost of the industry's fixed assets. *CC* is computed according to equation 4 by depreciating data from the Bureau of Economic Analysis (BEA [1987]) on industry investment expenditures for plant and equipment. The capacity depreciation schedules are based upon the BEA and reflect a 16-year straight-line schedule for equipment and a 30-year straight-line schedule for plant. The robustness of results to the assumed capacity depreciation schedule is assessed later. The *CC* model is applied separately to plant and equipment (P&E); also $(1\text{-}H)/(1\text{-}D)$ and $1\text{-}v\text{-}\tau(Z+Y)$ are computed separately for plant and for equipment. The entries in table 2 are weighted averages of the asset specific terms, where the weight is each asset's *CC* as a proportion of *CC* for combined P&E.

The industry market values rise throughout the sample period, although the increase is not monotonic nor do all industries move equiproportionately. Throughout the eighteen years, the average annual growth in *MV* equals 10.3 percent for Telephone Communications, 6.9 percent for Electric Services, and 4.0 percent for Natural Gas. Most of the increase in market values is due to increases in the asset base (and associated debt and equity issuances). *CC* rises by an average annual 12.0 percent, 11.0 percent, and 6.5 percent, in Telephone Communications, Electric Services, and Natural Gas, respectively. The asset base increases monotonically throughout the sample period for Telephone Communications and Electric Services, but for Natural Gas growth sputters in the 1980s. Indeed, for 1968 the asset base in Natural Gas (measured by *CC*) is about half as large as in Telephone Communications or Electric Services whereas in 1985 its asset base is less than one-quarter as large.

The *MV/CC* ratio is higher in 1968 than 1985 for each industry. For Telephone Communications the ratio ranges between 1.09 and 0.61 and its average is 0.74, implying the market value of financial liabilities for Telephone Communications averages 26 percent less than the current cost of the industry's fixed assets. For Electric Services the range for the ratio is between 1.24 and 0.45 and the ratio average is 0.70; for Natural Gas the range is 1.07 to 0.47 with an average of 0.69 . The ratio of market value to current cost for fixed assets exhibits substantial time series and cross-sectional variation. Typically, however, *CC* substantially overstates *MV*.

Table 2 presents evidence about the magnitude of the hypothesized bias embedded within the *CC* measurements. The total bias, denoted *B* and specified in equation 6, is partitioned into two components: $(1\text{-}H)/(1\text{-}D)$ and $1\text{-}v\text{-}\tau(Z+Y)$. These two terms are not related additively but they nonetheless represent the relative importance of the discounting and differential tax effects.

The term $(1\text{-}H)/(1\text{-}D)$ equals unity when the ratio of current-to-future pretax cash flow is the same for new assets as for used assets. Table 2 reveals that for Telephone Communications throughout the eighteen year sample period the term averages 0.7736 . This suggests that the ratio of current-to-future pretax cash flow is less for used assets than for new assets. Due to the discounting effect, current cost is hypothesized to exceed by about twenty-three percent the economic value of fixed assets.

The term $1\text{-}v\text{-}\tau(Z+Y)$ equals unity when the effective tax rates on income from new and used assets are equal. Table 2 reveals that for Telephone Communications this term averages 0.8796, suggesting the discounted tax liability is less for new assets than used assets. Due to the differential tax effect, current cost is hypothesized to exceed by about twelve percent the economic value of fixed assets.

The bias induced jointly by the discounting and differential tax effects is reflected through *B*; when *B* equals unity *CC* hypothetically equals the present value of expected after-tax cash flows. As shown in table 2, *B* averages 0.7292 for Telephone Communications. This suggests that *CC* overstates economic value by about 27 percent. *B* declines throughout the sample period, thereby implying an increase from 1968 to 1985 in the hypothetical bias embedded in the *CC* measurements. Similar trends are apparent for the other industries.

Further inspection of table 2 shows the likely source of decline in *B*. The bias induced by the discounting effect is fairly stable; $(1\text{-}H)/(1\text{-}D)$ fluctuates between 0.74 and 0.80 for Telephone Communications. The bias from the differential tax effect varies substantially; $1\text{-}v\text{-}\tau(Z+Y)$ drops from 0.96 to 0.81. Indeed, inspection of the first two columns reveals that from 1968 to 1985 the present value of tax depreciation deductions for new assets (*Z*) increases while, at the same time, the discounted deductions for used (in-place) assets (*Y*) decreases. The gap *Z-Y* for Telephone Communications opens from 5.39 cents to 26.83 cents, an increase of almost 400 percent. The decline in *B* and the accompanying hypothetical increase in the overstatement of economic value by *CC* is largely attributable to differential tax effects.

Estimation of the Empirical Model

The statistical significance of the hypothetical bias embedded within the *CC* measurements is assessed through estimation of equation 3:

$$\log MV_t = a_o + a_1 \log CC_t + a_2 \log B_t + e_t. \qquad (3)$$

The current cost model asserts that *CC* valuation measurements are unbiased estimates of economic values, in which case $a_1=1$ and $a_0=a_2=0$. The tax-adjusted fundamental value model, on the other hand, asserts that the *CC* measurements contain a systematic bias due to discounting and differential tax effects. The *FV* model enables measurement of *B* and posits that $a_1=a_2=1$ and $a_0=0$. The significance level on a_2 provides the basis for tests on the incremental information content of *B*.

Equation 3 is estimated for the annual time period 1968 through 1985 with the data constructed in the previous section. There are 18 times series and 3 cross sectional units, thereby yielding 54 observations for each variable *MV*, *CC*, and *B*. Results are shown in table 3.

The first row shows ordinary least squares (OLS) estimates for the case in which coefficients are allowed to vary between industries. In all cases the industry data are stacked and one ordinary least squares equation is estimated. Dummy variables are used, as appropriate, to allow estimated slope or intercept coefficients to vary between industries. The estimated intercepts (a_0) in all three industries are indistinguishable from zero. The coefficients (a_1) on log*CC* in all three industries are significantly different from zero but indistinguishable from unity. The results thus far support the predictions from both valuation models and offer support for the validity of the variables construction and the empirical estimation. The coefficients (a_2) on log*B*, predicted to equal zero by the *CC* model, are significantly different from zero as predicted by the *FV* model for Electric Services (SIC 491) and Natural Gas (SIC 492). For Telephone Communications (SIC 481) the coefficient a_2 is indistinguishable from zero.

Further analysis (not listed in table) shows that the qualitative results are robust to the assumed depreciation schedule. Reconstructing *B* and *CC* with accelerated depreciation schedules (200 percent and 100 percent declining balance for equipment and structures) and re-estimating row 1 yields coefficients (and t-statistics) in SICs 481, 491, and 492 for a_1 of 1.04 (10.4), 0.89 (9.1), and 1.00 (6.5); for a_2 industry coefficients are 2.21 (2.4), 2.75 (4.1), and 2.22 (3.9). Reconstructing *B* and *CC* with decelerated depreciation schedules (0.75 and 0.90 beta-decay for equipment and structures) yields estimates in SICs 481, 491, and 492 for a_1 of 0.97 (11.4), 0.82 (9.4), and 0.88 (6.5); for a_2 industry coefficients are 1.50 (2.1), 2.12 (3.9), and 1.73 (3.6). The estimated coefficients on log*FV* are significant in *all* industries for all alternative schedules.

Row 2 shows the results of enforcing an equality constraint across industries on a_2. As predicted by the *V* model, the estimated coefficient (2.52) is statistically different from zero (and unity) at the one percent significance level.[2] Furthermore, the statistical equality of a_2 across industries cannot be rejected.[3] Log*B* is a statistically significant explanatory variable for *MV*. This finding supports the hypothesis of this study that (i) discounting and differential tax effects systematically bias *CC* measurements of asset values, and (ii) the bias is measurable through reliance on the *FV* model.[4]

Two other restricted versions of equation 3 are estimated. Row 3 shows the results of an equality constraint on both a_1 and a_2. The coefficients on log*CC* and log*B* equal 0.98 and 2.35, respectively, and each is statistically distinguishable from zero at the one percent significance level.[5] Finally, row 4 examines the case in which a_0, a_1 and a_2 are the same in all three industries:

$$\underset{}{\log MV} = \underset{(2.9)}{0.78} + \underset{(37.3)}{0.97}\log CC + \underset{(10.6)}{2.32}\log B\,. \tag{7}$$

The t-statistics testing for zero-equality are presented in parentheses. All coefficients are significantly different from zero.[6]

Summary

According to both the *CC* and *FV* valuation models, the hypothesized coefficient (a_1) on log*CC* in equation 3 is unity. At 0.97, the estimated coefficient is indistinguishable from unity and significantly different from zero.[7] There is, as conventional wisdom maintains, a one-to-one (partial) correspondence between financial market value and the current cost of underlying assets.

According to the *FV* model, *CC* measurements are incomplete assessments of economic value due to the biasing influence of discounting and differential tax effects. The bias is measured in the *B* variable. The estimated coefficient (a_2) on log*B*, at 2.32, is highly significant and implies that *CC* measurements contain a systematic bias.

Valuation measurements computed as *CC*x *B* possess significant incremental information relative to measurements computed simply as *CC*. The *CC*x *B* measurement is the estimate of tax-adjusted fundamental value. Just as the *CC* model takes account of changing price levels ignored by the *HC* model, so does the *FV* model take account of discounting and tax effects ignored by the *CC* model. Under some conditions the *HC*, *CC*, and *FV* models collapse to the same specification and yield identical measurements. More generally, however, one model is an extension of the other, with the more sophisticated model incorporating incremental information ignored by the more primitive model.

Endnotes

1. Variants of the *FV* model are used for analyzing the effects of federal tax reform on stock prices by Downs and Hendershott [1987], and Downs and Tehranian [1988]. Downs [1992] uses the *FV* model to explain time-series and cross-sectional variation in industry Tobin's *Q* ratios. Downs and Shriver [1992] compare used asset transaction prices from secondary markets with the valuation estimates generated by the *HC*, *CC*, and *FV* models.

2. The magnitude of the coefficient, 2.52, exceeds its hypothesized value of unity. The coefficient is an estimate of the elasticity between *MV* and *B*. These results suggest that, whereas *MV* moves one-for-one with *CC*, the sensitivity of *MV* to *B* is greater than unity.

3. The sum-of-squared residuals (SSR) for the unrestricted equation (row 1) is 0.5737 and there are 9 estimated coefficients. The SSR for the restricted equation (row 2) equals 0.5870 and there are 7 estimated coefficients. A test statistic for the equality of a_2 across industries is [(.5870-.5737)/(9-7)]/[.5737/45], which equals 0.5216 . The F-statistic critical value at the five percent significance level is 3.21; the hypothesis for the equality of a_2 is *not* rejected.

4. The Durbin-Watson (DW) statistic for row 2, at 1.12, indicates the presence of residual autocorrelation. Estimation with a correction for first order serial correlation yields an equation qualitatively similar to the one in row 2. The estimated coefficient on log*B* equals 1.64 with a t-statistic of 3.3 and a DW of 1.54 .

5. The SSR for row 3 is 0.5820 and there are 5 estimated coefficients. The incremental SSR for row 3 relative to either rows 1 or 2 is such that the equality of a_2 across industries is *rejected* at the one percent significance level.

6. The joint hypothesis on the equality across industries of the coefficients a_0, a_1, and a_2 is rejected. Also, the DW of 0.80 indicates residual autocorrelation. Estimation of equation 7 with correction for first-order serial correlation yields coefficients on log*CC* and log*B* of 0.92 and 1.66 with t-statistics of 23.2 and 4.6, respectively, and a DW of 1.54 .

7. Estimation of equation 7 with the *B* and *CC* constructed with the aforementioned alternative capacity depreciation schedules yields coefficients (and t-statistics) for log*CC* and log*B* of 0.96 (41.3) and 2.39 (14.5) with accelerated depreciation and 0.96 (40.2) and 2.24 (13.9) with decelerated depreciation. The inferences are qualitatively the same.

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